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Abstract

We examine the response of real exchange rates to shocks in real exchange rate determinants, a monetary policy shock, and a fiscal policy shock in 30 countries over the period 1970-2008. The country set is divided into 4 groups - European, developed-country, Asian developing-country, and non Asian developing-country groups. We propose and apply a new approach, i.e. we employ a panel Bayesian structural vector error correction model, and we impose sign restrictions with a penalty-function approach to identify the shocks. We find that most of our impulse response analysis is in line with economic theories. Specifically, there is strong evidence that trade liberalization generates a real depreciation and an increase in government spending leads to a real appreciation over the long run. We also find that a contractionary monetary policy shock has only short-run impacts on real exchange rates, corresponding to the long-run neutrality of monetary policy. The responses to a productivity shock are interesting, i.e. productivity growth in traded sectors has no effect on the real exchange rate of the Asian developing-country group, and it leads to a long-run real appreciation in the non Asian developing-country group. In contrast, this shock causes a real depreciation in the European country group over the long run. Variance decomposition suggests that international trade policy contributes the most to real exchange rate movements in most country groups, with the exception of the non Asian developing-country group, for which fiscal policy via government spending seems to be the most important.

JEL Classification: C33, C51, E52, F31

Keywords: Real exchange rate, Vector error correction model, Monetary policy shocks, Sign restriction, Penalty function, Real exchange rate fundamentals, Identification

1 Introduction

The study of exchange rate behavior is motivated by lessons we learnt from the 1997 Asian Financial crisis. The crisis led to economic meltdown in many regional economies due to financial contagion. Many economists believe that the main cause of the crisis was a rapid change of exchange-rate regime from fixed exchange rates to a floating exchange in Thailand. This caused a dramatic depreciation of the Thai baht (and finally the

Thai baht reached its lowest point in January 1998), and then spread rapidly to put depreciation pressure on other foreign currencies in Asia; for example, in Malaysia, the ringgit plunged by 37.4 percent over the period from July 1 to September 30, 1997, and in South Korea, the Korean won dropped dramatically by more than 150 percent by the end of 1997. The collapse of foreign exchange markets exacerbated the failure in other financial markets, the breakdown of whole sectors, and finally economic downturn in many regional economies in Asia. From this situation, we learn that the adjustment process of foreign exchange markets can play an important role in financial crises, especially in a period of international financial integration.

The study of exchange rate movements and the relationship between the real exchange rate and its determinants have received considerable attention in the literature[1]. The initial works have simply used time series techniques such as Engle and Granger (1987) and Johansen (1988) cointegration tests[2]. This can lead to imprecise estimation and inconclusive hypothesis testing when spans of data are short.

More recent studies have turned to panel data cointegration methods. For instance, Chinn (1999) looks at the role of government consumption, the sectoral productivity differential, and the international terms of trade in determining the movement of real exchange rates using a panel of annual data in 14 OECD countries over the period 1970-1991. He constructs a variable of sectoral productivity in which traded sectors include agriculture, mining, manufacturing, and transportation, while nontraded sectors include all other services. Chinn estimates a panel error correction form using nonlinear least squares regression and concludes that an increase in relative traded sector productivity induces a long-run real appreciation while government spending and the terms of trade have no long-run effects on real exchange rate movements. These results are inconsistent with the work of Galstyan and Lane (2009) that investigate the relationship between government spending and the long-run real exchange rate for a panel of 19 advanced economies over 1980-2004. Using panel dynamic ordinary least squares (DOLS) estimation, they find that government consumption induces a long-run real appreciation[3].

In a related paper, Lee, Milesi-Ferretti, and Ricci (2008) examine the effects of economic fundamentals on the real effective exchange rate in a panel of 48 industrial countries and emerging markets over the period 1980-2004 (25 years). The underlying fundamentals include the productivity differential, net foreign assets, the commodity terms of trade, government consumption, and a trade restriction index. They pay more attention to the construction of a sectoral productivity variable, i.e. they rely on a six-sector classification - (i) agriculture, hunting, forestry, and fishing; (ii) mining, manufacturing, and utilities; and (iii) transport, stor-

[1] See, for example, Rogoff (1996), Edwards (1989), and Edwards and Savastano (2000).
[2] See, for example, Strauss (1995), Strauss (1996), and Kakkar (1996).
[3] This finding corresponds to the work of Froot and Rogoff (1996) and De Gregorio, Giovannini, and Wolf (1994).

age, and communication are classified as traded sectors, whereas (iv) construction; (v) wholesale and retail trade; and (vi) other services are nontraded sectors. The long-run relationships between the real exchange rate and the proposed set of real exchange rate fundamentals are estimated using DOLS methodology, developed by Stock and Watson (1993). They find that an increase in net foreign assets, the domestic productivity of tradables relative to nontradables, the commodity terms of trade, the extent of trade restrictions, and government consumption tend to appreciate a country's equilibrium exchange rate.

Although there has been a resurgence of interest in the relationship between the real exchange rate and its determinants, most of this work has focused on developed countries while the empirical evidence on this relationship for developing countries is not as abundant. Moreover, the construction of a productivity variable in traded and nontraded sectors is mostly based on arbitrary methods; that is, the classification approaches of industries also seem unreasonable as they use the same patterns of traded and nontraded sectors among industries in different countries. In addition, several existing papers examine the causal effects of exchange rate determinants on real exchange rates by relying only on the study of cointegrating relationships. However, the presence of a cointegrating relationship does not provide information on the direction of causality between the real exchange rate and its fundamentals.

The primary focus of this paper are to explore the role of a set of real exchange rate determinants derived from the Edwards (1989) and the Balassa-Samuelson models (Balassa 1964, Samuelson 1964), monetary policy, and fiscal policy in explaining real exchange rate behavior, and to assess quantitatively if these factors have played an important role in driving real exchange rate movements. We employ an unbalanced panel data set that includes annual time series from 1970 to 2008 for thirty countries (sixteen developed countries and fourteen developing countries). The country set is divided into four country groups: European, developed-country, Asian developing-country, and non Asian developing-country groups. We apply a novel approach for classifying traded and nontraded sectors, introduced in the first chapter, to construct variables of sectoral productivity and relative price of nontraded goods to traded goods. The long-run cointegrating relationship between the real exchange rate and its determinants is examined by using the Pedroni (2000) panel cointegration test together with the parsimonious approach of Davidson (1998). Panel DOLS estimation suggested by Mark and Sul (2003) is then employed to estimate the panel cointegrating vectors. The most innovative feature is that we use a panel Bayesian structural vector error correction model and impose sign restrictions with a penalty-function approach developed by Uhlig (2005) and Mountford and Uhlig (2009) in order to examine the effects of four shocks - productivity improvement, trade liberalization, contractionary monetary policy, and government spending - on real exchange rate movements. For all the country groups examined, impulse responses and variance decomposition are considered.

Our results show that there is a long-run cointegrating relationship between the real exchange rate and the sectoral productivity differential, real GDP growth, government consumption, the degree of openness in the economy in almost all country groups examined. The exception is the Asian developing-country group, in which the traded-nontraded productivity differential has no long-run cointegrating relationship with the real exchange rate. Most of our impulse response analysis is in line with economic theories. Firstly, our findings for all country groups examined confirm the traditional view that trade liberalization will generate a real depreciation over the long run. Secondly, as expected, higher government spending tends to induce a long-run real appreciation in almost all country groups. Thirdly, our analysis shows that a contractionary monetary policy shock has no long-run effect on real exchange rates in any of the country groups examined, even though for some country groups, the shock has a small influence on real exchange rates in the short run. This is consistent with the long-run neutrality of monetary policy.

In addition, the results show that a productivity shock causes different effects on real exchange rates in each country group. Specifically, we find that after a shock on productivity improvement in traded sectors, real exchange rates appreciate only in the non Asian developing-country group while the shock has no effect on real exchange rates of the Asian developing-country group. This finding is in line with the controversy regarding real exchange rate behavior among fast-growing countries in Asia. For developed countries, we find opposing evidence that as productivity growth in traded sectors increases, real exchange rates tend to depreciate[4]. One plausible explanation for this is that the shock may induce factor-augmenting progress in nontraded sectors, thereby leading to positive supply effects in nontraded goods. If this generates excess supply in nontraded goods and thus a fall in their prices, the shock may be followed by a real depreciation.

Furthermore, the variance decomposition shows that the traded-nontraded productivity differential, trade liberalization, monetary policy and government spending are important for explaining the dynamics of real exchange rates. Our results suggest that trade liberalization associated with international trade policy is the main determinant of real exchange rate movements in almost all country groups, with the exception of the non Asian developing-country group for which government spending associated with fiscal policy seems to be the most important.

The remainder of this paper is organized as follows. Section 2 provides the theoretical background and the reasons behind choosing each real exchange rate determinant in the model. Section 3 outlines the model and

[4] These results of a productivity shock correspond to those in the first chapter. That is, in the first chapter we found that productivity improvement in traded sectors relative to that in nontraded sectors leads to a real depreciation in developed countries while it causes a real appreciation in developing countries.

the econometric approaches used for identifying shocks in productivity improvement, trade liberalization, contractionary monetary policy, and government spending. Section 4 presents the data set, econometric methodology, and the empirical analysis. Section 5 concludes.

2 Theoretical Framework and Real Exchange Rate Determinants

Several empirical studies of real exchange rate behavior suggest that the real exchange rate might not just depend on the traded-nontraded productivity differential as mentioned in the Balassa Samuelson hypothesis (BSH). In this paper we extend the simple Balassa-Samuelson framework into a more general model of the real exchange rate and include more exchange rate determinants.

A number of models are developed to analyze the determination of the equilibrium real exchange rate. One of the first models of exchange rate determination is the monetary model which is a combination of the quantity theory of the demand for money and purchasing power parity (PPP). The monetary model describes the effect of shocks on a floating exchange rate and on the balance of payments based on the assumption of perfect flexibility of prices. However, it is completely inadequate for explaining real exchange rate determination, especially in the long run, due to the failure of PPP. Unlike the monetary model, the Mundell-Fleming (M-F) model (Mundell, 1962; Fleming, 1962) provides an analysis of the determination of the real exchange rate using a macroeconomic model, relying on the Keynesian tradition that a change in aggregate demand determines the level of economic activity, while aggregate supply takes the role of fixing the price level. It allows for the absence of PPP, by allowing imperfect capital mobility and static expectations. The definition of the equilibrium real exchange rate in this model is the relative price of international to domestic goods that leads to simultaneous equilibrium in the money market, the domestic goods market, and the international goods market. The main weakness of this model, however, is that it is restricted by the assumptions of a constant price level and static expectations, thereby limiting its relevance to only the very shortest horizons.

Dornbusch (1974, 1980) provides a hybrid model of exchange rate determination based on the above two models. In the short run, the Dornbusch model relies on the stickiness of prices in the product and labour markets as in the M-F model, while its long-run characteristics are aligned with the monetary model - the model is based on the assumptions of complete price flexibility and full employment in the long run. Dornbusch defines the equilibrium real exchange rate as the relative price of tradables to nontradables that simultaneously equilibrates the markets of tradable and nontradable goods. However, the weakness of this model is that it does not allow for changes in the terms of trade between tradable goods to have an influence on the equilibrium real exchange rate.

Several models of real exchange rate determination, including the models mentioned above, have the main limitation that they do not allow for the different effects of temporary and permanent changes in the exchange rate determinants. To rectify this, the long-run real exchange rate framework developed by Edwards (1989) is used extensively in the study of real exchange rate behavior, especially in developing countries[5]. The Edwards model is an intertemporal duality model, and it defines the equilibrium real exchange rate as the relative price of traded to nontraded goods which leads to simultaneous internal and external equilibria of the economy, given long-run equilibrium values of other relevant variables. In this model, internal balance means that the nontraded market currently clears and will be in equilibrium in the future, thus implying the existence of unemployment at the natural level at the equilibrium. External equilibrium means that the current account balances in the current period and in the future will satisfy the intertemporal budget constraint i.e. the discounted sum of the current account has to be zero, and thus it is compatible with sustainable capital flow in the long run. The Edwards model also shows that the long-term real exchange rate relies on real variables only, while both monetary and real variables can influence the real exchange rate in the short run. In contrast to PPP, the equilibrium real exchange rate can change when there are changes in long-term real exchange rate determinants. Since the focus of this study is on the long-term behavior of real exchange rates, especially in developing countries, we will develop our empirical model including the main long-run real exchange rate determinants called real exchange rate fundamentals, as suggested by Edwards (1989).

There are two main categories of real exchange rate fundamentals in the Edwards model. The first category consists of external fundamentals, and it includes the international terms of trade and real interest rate differentials. The second category, called domestic fundamentals, can be distinguished into two groups - domestic policy-related fundamentals, which include trade barriers, taxation, capital controls, and government consumption; and the domestic nonpolicy-related fundamental, which is technological progress as measured by the growth rate of real Gross Domestic Product (GDP). There are several previous studies on real exchange rate misalignment that rely on the Edwards model. These studies include Aron et al. (1997), Elbadawi and Soto (1997), and Ghura and Grennes (1993), and they find evidence that the terms of trade, the degree of openness in the economy, public expenditure, and capital flow are real exchange rate determinants.

According to the BSH, long-term real exchange rate movements involve two main channels: the first

[5] See, for example, Mkenda (2001), Elbadawi (1994), and Krumm (1993).

channel links the real exchange rate to the relative price of nontraded goods between domestic and foreign countries; and the second channel relates the relative prices of nontraded goods and traded-nontraded productivity differentials. Moreover, the Balassa-Samuelson prediction is based on the assumption of a positive relationship between the traded-nontraded productivity differential and economic growth. Therefore, the real exchange rate determinants in the Balassa-Samuelson model involve the relative price of nontraded goods to traded goods, the traded-nontraded productivity differential, and economic growth at home and abroad.

Combining the fundamentals associated with the BSH framework and the Edwards model, the real exchange rate fundamentals are the traded-nontraded productivity differential, the relative price of nontraded goods, the growth rate of real GDP, the degree of openness in the economy, the international terms of trade, capital flows, and government consumption. These are also the set of potential fundamentals found in the extensive existing literature on the determinants of the long-run equilibrium real exchange rate. Therefore we will develop our model by including these real exchange rate determinants. In the following, we will discuss traditional views and theoretical concepts associated with the BSH and the Edwards Frameworks.

2.1 The Balassa-Samuelson Model

2.1.1 Productivity gain, relative price of nontraded goods, and economic growth

One of basic theories of long-run real exchange rate behavior is the BSH. The BSH provides an explanation of long-run real exchange rate behavior based on productivity differentials between traded and nontraded goods. The explanation is that a higher productivity differential in the traded sector relative to the nontraded sector leads to higher production costs for traded goods, and due to labour mobility, an increase in productivity of the traded sector raises the wage in the nontraded sector. A higher price of nontraded goods is required to maintain profitability in the nontraded sector. This is the first main channel of the BSH that represents the relationship between the traded-nontraded productivity differential and the relative price of nontraded goods. Then the implication of a higher relative price of nontraded goods to traded goods at home relative to that of other countries implies an appreciation of the real exchange rate. This is the second main channel of the BSH.

Overall, the real exchange rate appreciates when productivity gains in the traded sector are higher in the domestic country than in other countries. As productivity gains in the traded sector are assumed to be mainly responsible for economic growth, this leads to the well-known prediction that faster economic growth should be accompanied by a real appreciation. Therefore we can conclude that the BSH relies on three main long-run real exchange rate fundamentals: the price of nontraded goods relative to traded goods, the traded-nontraded productivity differential, and the relative growth rate of real GDP between the domestic and foreign countries. From this, these three fundamentals are considered in our empirical model.

2.2 The Edwards Model

We summarize the basic background of the Edwards model, before discussing the importance of each real exchange rate fundamental in detail. Edwards (1989) explains the equilibrium path of real exchange rates using duality theory and an intertemporal general equilibrium model of a small open economy. The economy consists of optimizing consumers and producers as well as a government with three goods: exportables (X), importables (M), and nontradables (N). The factors of production are capital, labour, and natural resources. The benchmark framework starts with the assumptions of full employment, no price rigidities, no international credit rationing, perfect foresight, perfect competition, and a constant labour force. These assumptions are relaxed later.

According to the Edwards (1989) model, there are two periods: the current (period 1) and the future (period 2). Note that throughout this section, no tilde and a tilde ($\tilde{}$) over a variable denote a variable in period 1 and period 2, respectively. The superscript * denotes a foreign variable and no superscript implies a domestic variable. We let R be the revenue function and E be the expenditure function, and R and E subscripts indicate partial derivatives with respect to an argument (e.g., R_s refers to the partial derivative of R with respect to s, or R_{ps} refers to the second derivative of R with respect to p and s). Residents in this small country can borrow and lend internationally but they are subject to a nonprohibitive tax, implying that the domestic real interest rate is higher than the world interest rate. Debts are repaid in period 2. Moreover, tariffs are used for import commodities in both periods. The model emphasizes real variables - there is no money, nor any nominal variable - and the world price of exportables is used as the numeraire.

We start with the firm's problem. Firms produce goods using constant returns to scale technology under perfect competition, and they satisfy the intertemporal budget constraint. In each period, the production side is characterized by revenue functions, R and $\widetilde{R}$. The optimizing firms produce X, M, and N by maximizing their revenue functions subject to prevailing domestic prices, technology and the available factors of production. For period 1, the revenue function is given by

$$R = \max\{Q_X + pQ_M + sQ_N | F(Q, V) \leq 0\} \quad or \quad R = R(p, s, V). \tag{1}$$

Following conventional notation, Q_X, Q_M, and Q_N are the quantities of exportables, importables, and nontradables produced in period 1. Q is a vector of quantities produced, V is a vector of factors of production, and $F(\ )$ is the production possibilities function. The domestic prices of importables and nontradables in terms of exportables in period 1 are p and s, respectively. This function gives the maximized value of output in period 1 in terms of exportables. The revenue function in period 2 can be formed in a similar way.

The supply function for each good can be derived from the first derivative of the revenue function with respect to the price of that good. For example,

$$\frac{\partial R}{\partial p} = R_p = Q_m(p, ...) \quad \text{is the supply function for } M \text{ in period 1.}$$

The convexity of the revenue function implies that $R_{pp} = \frac{\partial Q_M}{\partial p} \geq 0$. That is, there is an upward slope in the supply function. Suppose that there are no intermediate inputs and each good competes for the given amount of resources. This implies that the cross-price derivatives of revenue functions are negative, i.e. $R_{ps} = \frac{\partial Q_M}{\partial s} \leq 0$.

Turning to the representative consumer's problem, Edwards assumes that the utility function is time separable, and each subutility function is homothetic. Each consumer consumes all three goods and maximizes the present value of utility, subject to the intertemporal budget constraint. That is, the consumer solves

$$max \ \ W\{U(C_N, C_M, C_X), \widetilde{U}(\widetilde{C}_N, \widetilde{C}_M, \widetilde{C}_X)\} \tag{2}$$

subject to

$$C_X + pC_M + sC_N + \delta(\widetilde{C}_X + \widetilde{p}\widetilde{C}_M + \widetilde{s}\widetilde{C}_N) \leq Wealth,$$

where W is the utility function, U and $\widetilde{U}$ are subutility functions in periods 1 and 2, C_N, C_M, C_X are the consumption of N, M, X, and δ indicates the domestic discount factor which is equal to $(1+r)^{-1}$where r is the domestic real interest rate in terms of exportables. $Wealth$ in the budget constraint refers to the discounted sum of the consumer's income from labour services rendered to firms, from the renting of capital stock to domestic firms, and from government transfers in periods 1 and 2.

Edwards derives the demand side of the model from a twice-differentiable concave expenditure function, E. This provides the minimum discounted value of expenditure required to obtain a level of utility $\overline{W}$ for given domestic prices in period 1 and 2, and the optimization problem is then written as follows:

$$E = \min\{C_X + pC_M + sC_N + \delta(\widetilde{C}_X + \widetilde{p}\widetilde{C}_M + \widetilde{s}\widetilde{C}_N)\} \tag{3}$$

subject to

$$W(U, \widetilde{U}) \geq \overline{W} \ .$$

It is obvious that equation (3) is a function of prices and utility, so it can be also written as

$$E = E\{\pi(p,s), \delta\widetilde{\pi}(\widetilde{p},\widetilde{s}); W\}, \tag{4}$$

where $\pi(\ )$ and $\widetilde{\pi}(\ )$ are exact price indices that can be viewed as unit expenditure functions for each period. Similar to the derivation of the supply function, the compensated (Hicksian) demand function for each good can be derived from the partial derivative of the expenditure function with respect to the price of that good. That is,

$$\begin{aligned}
E_p &= \frac{\partial E}{\partial \pi}\frac{\partial \pi}{\partial p} = E_\pi \pi_p = D_M(p, ...), \\
E_s &= \frac{\partial E}{\partial \pi}\frac{\partial \pi}{\partial s} = E_\pi \pi_s = D_N(s, ...), \\
E_{\widetilde{p}} &= \frac{\partial E}{\partial \widetilde{\pi}}\frac{\partial \widetilde{\pi}}{\partial \widetilde{p}} = E_{\widetilde{\pi}} \widetilde{\pi}_{\widetilde{p}} = \widetilde{D}_M(\widetilde{p}, ...), \quad \text{and} \\
E_{\widetilde{s}} &= \frac{\partial E}{\partial \widetilde{\pi}}\frac{\partial \widetilde{\pi}}{\partial \widetilde{s}} = E_{\widetilde{\pi}} \widetilde{\pi}_{\widetilde{s}} = \widetilde{D}_N(\widetilde{s}, ...),
\end{aligned}$$

where D_M ($\widetilde{D}_M$) and D_N ($\widetilde{D}_N$) are the Hicksian demand functions for M and N in period 1 (2). $\pi_p(\widetilde{\pi}_{\widetilde{p}})$ and $\pi_s(\widetilde{\pi}_{\widetilde{s}})$ can be regarded as expenditure shares of M and N in period 1 (2).

The demand curve has a downward slope due to the concavity of the expenditure function that ensures that $E_{pp}, E_{ss}, E_{\widetilde{p}\widetilde{p}}$, and $E_{\widetilde{s}\widetilde{s}} < 0$. Expenditures in periods 1 and 2 can be substitutes since we assume a time separable utility function. That is, all intertemporal cross elasticities are positive. However, in each period, two goods can be complements (or substitutes) so that the intratemporal cross elasticities can be negative (or positive).

The government is another agent in the model. The government consumes tradable and nontradable goods, and its expenditure is financed from four sources: nondistortionary taxes, import tariffs, the taxation of foreign borrowing by the private sector, and borrowing from abroad. As in the case of the private sector, the government has to satisfy its intertemporal budget constraint.

According to the Edwards model, the following equations describe the economy at external and internal equilibria:

$$R(1,p,s,V,K) + \delta\widetilde{R}(1,\widetilde{p},\widetilde{s},\widetilde{V},K+I) - I(\delta) - T - \delta\widetilde{T} = E\{\pi(1,p,s), \delta\widetilde{\pi}(1,\widetilde{p},\widetilde{s}), W\}, \tag{5}$$

$$G_X + p^* G_M + sG_N + \delta^*(\widetilde{G}_X + \widetilde{p}^*\widetilde{G}_M + \widetilde{s}\widetilde{G}_N) = \tau(E_p - R_p) + \delta^*\widetilde{\tau}(E_{\widetilde{p}} - \widetilde{R}_{\widetilde{p}}) + b(NCA) + T + \delta^*\widetilde{T}, \tag{6}$$

$$R_s = E_s + G_N, \quad \widetilde{R}_{\widetilde{s}} = E_{\widetilde{s}} + \widetilde{G}_N, \tag{7}$$

$$p = p^* + \tau, \quad \widetilde{p} = \widetilde{p}^* + \widetilde{\tau}, \tag{8}$$

$$\delta \widetilde{R}_K = 1, \tag{9}$$

$$P_T^* = \gamma P_M^* + (1-\gamma) P_X^*, \quad \widetilde{P}_T^* = \gamma \widetilde{P}_M^* + (1-\gamma) \widetilde{P}_X^* \quad \text{where} \quad P_X^* = \widetilde{P}_X^* = 1, \tag{10}$$

$$RER = \frac{P_T^*}{P_N}, \quad \widetilde{RER} = \frac{\widetilde{P}_T^*}{\widetilde{P}_N}, \quad \text{and} \tag{11}$$

$$RER = E \frac{(P_N^*)^{\theta^*} (P_T^*)^{1-\theta^*}}{(P_N)^{\theta} (P_T)^{1-\theta}} \propto \frac{(P_N^*)^{\theta^*} (P_T^*)^{1-\theta^*}}{(P_N)^{\theta} (P_T)^{1-\theta}}. \tag{12}$$

Equations (5)-(12) use the same notation as before, as well as additional new notation. For the latter, T is the lump sum tax, K is the capital stock, P_T is the world price of tradables, RER is the real exchange rate, $I(\ )$ is investment in period 1, τ is an import tariff, and G_X, G_M, and G_N are the quantities of government consumption in X, M, and N, respectively. Moreover, NCA is the private sector current account surplus in period 2 which is equal to $(\widetilde{R} - \widetilde{\pi} E_{\widetilde{\pi}})$, and $b(NCA)$ is the discounted tax value where the private sector has to pay for foreign borrowing. As before, no tilde and a tilde ($\widetilde{\ }$) over a variable denote a variable in period 1 and period 2 respectively, and E or R subscripted by a price variable indicates a partial derivative with respect to that variable. Note that the domestic discount factor, $\delta = (1+r)^{-1}$, is smaller than the world discount factor, $\delta^* = (1+r^*)^{-1}$, as the real domestic interest rate, r, is higher than the world real interest rate, r^*, due to a tax on foreign borrowing.

Equation (5) is the intertemporal budget constraint for the private sector. It says that the present value of income valued at domestic prices is equal to the present value of private expenditure. Equation (6) refers to the government intertemporal budget constraint, which implies that the discounted value of government expenditure is equal to the discounted value of government income from taxation consisting of the discounted sum of import tariffs, taxes on foreign borrowing paid by the private sector, and lump sum taxes in both periods.

Equation (7) represents the equilibrium conditions of nontradable goods in both periods. Equation (8) indicates the relation between the world price and the price of importables in the presence of import tariffs.

Equation (9) describes investment decisions that profit maximizing firms will add to the capital stock until Tobin's "q" equals 1. Equation (10) is the definition of the price index for tradables, where γ is the share of importables. Equations (11) provide definitions of the real exchange rate in the Edwards model. We add equation (12) into the Edwards model for the purpose of comparison. This equation is the definition of the real exchange rate used in our empirical model. Edwards defines the real exchange rate index as the ratio of the world price of tradables to the domestic price of nontradables, while the real exchange rate in our empirical study is based on the PPP condition and so it is measured as the proportion of the world price of a GDP bundle to the domestic price of the corresponding bundle. However, it is readily clear that due to the small open economy assumption, P_T, P_T^*, and P_N^* are determined abroad and they are therefore exogenous, implying that the real exchange rate in both definitions depends on the same single factor i.e. the domestic price of nontradables, P_N. Therefore, although our definition of the real exchange rate is different from that of the Edwards model, this does not change the qualitative analysis of the effects of a shock to real exchange rate behavior, and we still obtain the same conclusion as in the Edwards model. From this, we describe and summarize the Edwards model using his original definition of the real exchange rate.

In contrast to PPP, there is not a single equilibrium value of the real exchange rate in the Edwards model, but a vector of equilibrium real exchange rates $\mathbf{RER} = (RER, \widetilde{RER})$, which consists of real exchange rates that satisfy equations (5)-(11), given equilibrium values of other relevant variables. By analogy, as our study is based on the Edwards framework, the vector of equilibrium real exchange rates in our study consists of real exchange rates that satisfy equations (5)-(10) and (12). Overall, the vector of the equilibrium real exchange rates in Edwards model and our empirical representation of this model can be written as functions of the sustainable levels of all exogenous variables in the system as follows:

$RER = h(p^*, \widetilde{p}^*, \tau, \widetilde{\tau}, \delta, \delta^*, V, T, \widetilde{T}, G_X, \widetilde{G}_X)$,

$\widetilde{RER} = \widetilde{h}(p^*, \widetilde{p}^*, \tau, \widetilde{\tau}, \delta, \delta^*, \widetilde{V}, T, \widetilde{T}, G_X, \widetilde{G}_X)$.

The following provides details on the reactions of equilibrium real exchange rates to a change in real exchange rate fundamentals, as stated in traditional views and the Edwards model. This outlines the reasons behind the selection of each variable in our empirical model. Edwards analyzes the effects of a shock by using a simplified version of a general model of equilibrium real exchange rates that accounts for all essential aspects of the question we are addressing. Note that we simply provide a descriptive summary of some analysis that is associated with Edwards model. We refer the reader to the formal mathematical and diagrammatic analysis in Edwards (1989), for more details.

2.2.1 Technological Progress

As mentioned above, the conventional understanding of the effects of a technological change on real exchange rates is derived from the Balassa-Samuelson theory. The argument is that a country with a higher productivity level in traded goods relative to nontraded goods than that of other countries will create a higher relative price of nontraded goods, implying an appreciation of the real exchange rate.

According to the Edwards model, the effects of technological progress can be considered using the above general model by adding a shift parameter into the revenue functions for both periods (See Edwards (1989) for details). The type of progress - product augmenting or factor augmenting - and the rate at which productivity improves across different sectors will have different effects on the real exchange rate. For all types of shocks to productivity gains, an income effect will generate higher demand for everything and hence an increase in the price of nontradables which, in turn, will induce a real appreciation in both periods. In addition to this effect, technology gains will create supply effects. If the progress is of the factor augmenting type, the Rybczynski (1955) principle applies. That is, an increase in a factor of production causes more than a proportional increase in goods that are intensive in this factor, and a fall in other goods. Therefore, if a factor of production which is intensively used in the production of nontradables increases, this will lead to an increase in nontradables relative to others and a decline in their price, thus inducing a real depreciation. Moreover, if the product augmenting technological improvement increases the quantity of nontradables until there is excess supply, then the price of nontradables has to decrease to restore equilibrium. This situation will lead to a depreciation of the real exchange rate. Therefore if the supply effects dominate the demand effects in response to a technological shock, then it is possible that technological progress can cause a real depreciation. This analysis leads to a different conclusion from the BSH because according to the BSH, a productivity gain in tradables relative to nontradables will cause a real appreciation. In the Edwards model, it can cause either a real depreciation or a real appreciation depending on the domination of either supply or demand effects.

2.2.2 Trade liberalization

The traditional view is that a trade liberalization (trade restriction) generates a long-run real depreciation (appreciation). The common argument is that trade liberalization achieved by decreasing import tariffs implies a lower domestic price of importable goods, and hence a rise in the demand for these goods. This will cause a trade account deficit. The real exchange rate has to depreciate in order to restore external equilibrium.

The effect of trade liberalization in the Edwards model is a priori indefinite, depending on the assumptions imposed. Edwards uses a simplified version of the general model above to analyze the response to this shock. He assumes that there is no government consumption and no investment. He also assumes perfect access to the world capital market i.e. no taxes on foreign borrowing. This ensures that the domestic and world discount factors are equal ($\delta = \delta^*$). The simplified model can be characterized by the equations below:

$$R(1, p, s; V) + \delta^* \widetilde{R}(1, \widetilde{p}, \widetilde{s}, \widetilde{V}) + \tau(E_p - R_p) + \delta^* \widetilde{\tau}(E_{\widetilde{p}} - \widetilde{R}_{\widetilde{p}}) = E[\pi(1, p, s), \delta^* \widetilde{\pi}(1, \widetilde{p}, \widetilde{s}), W], \tag{13}$$

$$R_s = E_s, \quad \widetilde{R}_{\widetilde{s}} = E_{\widetilde{s}}, \tag{14}$$

$$p = p^* + \tau, \quad \widetilde{p} = \widetilde{p}^* + \widetilde{\tau}, \tag{15}$$

$$CA = R + \tau(E_p - R_p) - \pi E_\pi, \quad \text{and} \tag{16}$$

$$RER = \frac{P_T^*}{P_N}, \quad \widetilde{RER} = \frac{\widetilde{P}_T^*}{\widetilde{P}_N}. \tag{17}$$

Equation (13) represents the intertemporal budget constraint. Equation (14) provides the equilibrium conditions for the market of nontradables in periods 1 and 2. Equation (15) describes the relationship between the domestic and world import prices, which is the same as equation (8) in the general model. Equation (16) defines the current account in period 1 to be the difference between income and total expenditure in period 1. Equations (17) provide the definition of the real exchange rate in the Edwards model.

For the sake of clarity, Edwards supposes that agents (correctly) expect that the government will impose an import tariff in period 2. First, he assumes no import tariffs at the beginning in each period but he later relaxes this assumption. Consider the effects on market equilibrium in period 1, first. With the tariff, consumers will expect that consumption in period 2 will be relatively more expensive. Intertemporal substitution in consumption will imply that they will consume more in period 1, thereby raising the demand for all goods and the overall price level including the price of nontradables in period 1. The magnitude of the increase in the price of nontradables will depend on the intertemporal degree of substitutability.

The imposition of an anticipated import tariff in period 2 will cause both intertemporal and intratemporal effects in this period. The intertemporal effect will be less consumption in period 2. This will lead to a reduction in the price of nontradables in period 2. For the intratemporal effect, the effect will depend on whether nontradables and importables are substitutes or complements in consumption. If they are substitutes, the imposition of the import tariff will increase the demand for nontradable goods and thus their price. In contrast, if they are complements, the price of nontradable goods will fall. The more plausible assumption is that they are substitutes and we will rely on this assumption in our analysis. Consequently, under the condition of net substitutability, the expected import tariffs in period 2 will tend to raise the price of nontradables and hence cause real appreciation in both periods.

In reality, the initial condition of no import tariff at the beginning is implausible, as most countries already have tariffs or other types of import barriers. With the relaxation of this assumption, the first-order income effects have to be considered. Suppose that there is an increased import tariff in period 1. This will lead to a negative income effect in both periods, implying a reduction in the demand for these goods and their price. Thus, the income effect will tend to depreciate the real exchange rate. In contrast, if all goods are substitutes in consumption, an increase in import prices will raise the price of nontradables and induce a real appreciation. Therefore, if the substitution effect dominates the income effect, an increase in tariffs in period 1 will lead to a real exchange rate appreciation in both periods, consistent with the traditional view.

For the empirical work below, the ratio of the sum of exports and imports to GDP, which is a common indicator of the degree of openness in the economy, is used as a proxy variable for trade liberalization.

2.2.3 Terms of Trade

The traditional view is that an improvement in the terms of trade will induce a real appreciation. That is, an increase in the world price of exportables will induce higher real income and hence higher demand for nontradables. Moreover, wages in the nontraded sector have to increase in order to reestablish equilibrium because the factors of production from other sectors are absorbed to produce tradable goods. Thus this generates a higher relative price of nontradables which, in turn, leads to an appreciation of the real exchange rate.

However, the substitution effect of a terms of trade shock should also be considered in addition to the income effect. In his analysis of the terms of trade, Edwards relies on the simplified version of the general model used for the study of trade liberalization above. To illustrate, let us consider the case of a permanent terms of trade improvement in periods 1 and 2. Under the assumption of net substitutability between importables and nontradables, a decline in the price of importables will induce higher demands for such goods, and lower the demand for nontradables. The price of nontradables then has to decrease in order to maintain equilibrium. This effect works in an opposite direction from the income effect discussed above. Therefore the response of the long-run real exchange rate to a terms of trade shock in the Edwards model depends on whether the intertemporal substitution or income effect dominates.

In our study, the terms of trade is defined as the ratio of the export price index to the import price index.

2.2.4 Capital Control

In general, a relaxation of capital controls via a decline in a tax on foreign borrowing will lead to capital inflow. As a result, expenditure in this country can exceed income, generating higher demand for all goods including nontradables. The price of nontradables has to increase to maintain equilibrium and this, in turn, will induce a real appreciation.

In the Edwards Model, the capital controls are captured by a tax on foreign borrowing, making the domestic interest rate higher than the world interest rate. Edwards simplifies his general model by ignoring unimportant aspects of a change in capital control. Thus, he assumes that no import tariff is used, and that the international prices of exportables and importables are constant and can be aggregated into a composite tradable good. As before, he further assumes that there is no investment and no government consumption so that all taxes are returned to the public. According to Edwards (1989), the simplified model can be written as below:

$$R(1, f; V) + \delta\widetilde{R}(1, \widetilde{f}; \widetilde{V}) + b(NCA) = E[\pi(1, f), \delta\widetilde{\pi}(1, \widetilde{f}), W], \tag{18}$$

$$b = (\delta^* - \delta) > 0, \tag{19}$$

$$R_f = E_f, \quad \widetilde{R}_{\widetilde{f}} = E_{\widetilde{f}}, \quad \text{and} \tag{20}$$

$$RER = 1/f, \quad \widetilde{RER} = 1/\widetilde{f} \quad \text{where } f = P_N/P_T^* \ . \tag{21}$$

Equation (18) is the budget constraint for the private sector and equation (20) is the equilibrium condition for nontradables in periods 1 and 2. Notice that in contrast to equation (13) above, the domestic and world discount factors are different in this model.

To illustrate the effects of a change in capital control, consider the case of a liberalization of the capital account. This can be caused by a reduction in the domestic interest rate. The reduction will induce both an intertemporal substitution effect and a welfare effect. For the intertemporal substitution effect, a reduction in the domestic interest rate is equivalent to an increase in the discount factor, so all goods in period 2 will become relatively more expensive than those in period 1. As a result, the consumers will substitute away from consumption in period 2 and consume more in period 1, inducing an increase in the demand for all goods. The price of nontraded goods will increase and this will generate an appreciation of the real exchange rate. The income effect will take place in addition to the intertemporal substitution effect. The relaxation of barriers to foreign borrowing will lead to a positive welfare effect. Thus, higher demand for everything will be followed by an increase in the price of nontradables, implying a real appreciation. According to the Edwards model, a reduction in a tax on foreign borrowing in period 1 leads to a real appreciation in period 1, consistent with the traditional view. However, the reaction in period 2 to the liberalization in capital control depends on whether the intertemporal substitution or the income effect dominates.

In the work that follows, real interest rate differentials at home and abroad are used as proxies for capital controls. In particular, a decrease in the difference between the domestic and world interest rates indicates the relaxation of barriers.

2.2.5 Government Expenditure

The common view is that an expansion in fiscal policy via an increase in government expenditure will generate higher demand for nontradable goods, raise their price, and thereby induce an appreciation of the real exchange rate.

It is worth noting that the traditional view disregards the composition of government expenditure (into nontraded sectors or traded sectors) and ignores the effects of financing the debt. Edwards shows that the impact of government expenditure on the real exchange rate depends on the allocation of expenditure across goods and the type of taxes used. To make his exposition clear, he assumes no import tariffs, no change in the terms of trade, no investment, and that the government does not face an inflation tax. The model is then given by

$$R(1, f, V) + \delta^* \widetilde{R}(1, \widetilde{f}, \widetilde{V}) - T - \delta^* \widetilde{T} = E(\pi(1, f), \delta^* \widetilde{\pi}(1, \widetilde{f}), W), \tag{22}$$

$$G_T + f G_N + \delta^* (\widetilde{G}_T + \widetilde{f} \widetilde{G}_N) = T + \delta^* \widetilde{T}, \tag{23}$$

$$R_f = E_f + G_N, \quad \widetilde{R}_{\widetilde{f}} = E_{\widetilde{f}} + \widetilde{G}_N, \quad \text{and} \tag{24}$$

$$RER = (1/f), \quad \widetilde{RER} = (1/\widetilde{f}). \tag{25}$$

Equation (22) is the budget constraint for the private sector. Equation (23) indicates the government's budget constraint. Equation (24) is the equilibrium condition for nontradables in periods 1 and 2.

Notice that contrary to the traditional approach, the government can consume both tradable and nontradable goods. Edwards separates the government's budget constraint into the two equations below:

$$(G_T + fG_N) - T = D, \quad \text{and} \tag{26}$$

$$D + \delta^*(\widetilde{G}_T + \widetilde{f}\widetilde{G}_N) = \delta^*\widetilde{T}. \tag{27}$$

From equations (26) and (27), it is clear that government expenditure which exceeds tax income (D) is financed by borrowing from abroad. This also shows that the discounted value of tax revenues must be enough to cover government debt and consumption in period 2.

The following two examples illustrate that a change in government expenditure can have an ambiguous effect on the exchange rate. First, assume that government spends more on nontradables in period 1. The excess of expenditure over revenues is financed by public debt and must be paid back in period 2. According to the traditional view, an increase in government consumption of nontradables in period 1 will create higher demand and thus a rise in the price of nontradables, generating a real appreciation. However, to finance the debt in period 2, the government will need to increase taxes and hence decrease household income. This will tend to reduce consumption in both periods and thus cause a real depreciation. Given these two channels, the effect of an increase in government spending on nontradables is a priori indefinite, depending on the sum of these two effects. However, in most cases, the former effect is dominant and the real exchange rate will appreciate under this condition.

Secondly, suppose that the government temporarily increases consumption of tradable goods, instead of nontradable goods. It is readily shown that a temporary increase in government consumption of tradables will cause a real depreciation in periods 1 and 2. However, it is more plausible that the consumption of government will focus more on nontradable goods.

Following the conventional approach, we use the share of government consumption in GDP as a measure of government demand for nontradable goods.

3 The Model and Identification Procedure

Our model contains domestic variables and foreign variables. Each individual country-group model is estimated separately. The key assumption for the purposes of estimation and inference is that foreign variables are weakly exogenous, compatible with a limited degree of weak dependence across idiosyncratic shocks.

Let us illustrate how to construct country-specific foreign variables. Throughout this section, all variables are in logarithmic form. Also, the superscript * denotes a foreign variable and no superscript indicates a domestic variable. We consider N+1 countries in the data set where country 0, the U.S., is used as the reference country. Instead of using the same foreign variables in each country, we use country-specific foreign variables, constructed by using country-specific weighted averages of domestic variables of all other countries or regions. In our study, country-specific weights are based on average trade shares over the period 2002-2008. The foreign variable is given by

$$g_{it}^{*} = \sum_{j=0}^{N} w_{ij} g_{jt}, \quad i = 0, 1, 2..., N,$$

where w_{ij} is the trade share of country j in the total trade (imports + exports) of country i, such that $w_{ii} = 0$ and $\sum_{j=0}^{N} w_{ij} = 1$. It is clear that the trade weights are used to capture the relative importance of country j to country i, thereby, generating foreign variables which are more appropriate than traditional foreign variables, i.e. U.S. series. Table B1 in Appendix B provides a 26×26 matrix of the trade shares used to construct the country-specific foreign variables in our analysis[6].

Our model uses real effective exchange rates, defined following Dees et al. (2007) as

$$\begin{aligned} q_{it} &= \sum_{j=0}^{N} w_{ij}(e_{it} - e_{jt}) + p_{it}^{*} - p_{it} \\ &= \overleftrightarrow{e}_{it} - \overleftrightarrow{e}_{it}^{*}, \end{aligned}$$

where $\overleftrightarrow{e}_{it} = e_{it} - p_{it}$, and $\overleftrightarrow{e}_{it}^{*} = e_{it}^{*} - p_{it}^{*}$ for $i = 0, 1, 2, ..., N$. Thus q_{it} is the real effective exchange rate, e_{it} is the nominal exchange rate with respect to U.S., and p_{it} is the consumer price index (CPI).

As our focus of this study is on the movement of real exchange rates in response to shocks in exchange rate fundamentals and policy applications, we include real exchange rates, their fundamentals, nominal interest rates, the inflation rate, and oil prices in each country-group model. Given the nature of macroeconomic variables and the assumption of a small open economy, we treat all domestic variables as endogenous.

In the following subsections, we describe the model and econometric methodology for the identification of shocks. The first subsection presents the model. The identification strategy, based on sign restrictions with a penalty-function approach is discussed in the second subsection. The third subsection provides the identifying assumptions and the implementation of this identification strategy.

[6] Five countries in the euro area are treated as a single economy.

3.1 Structural Vector Autoregression Model

We will outline the basic framework for constructing the model. We construct an individual model for each country group instead of an individual model for each country, so that a panel-estimation technique will improve the efficiency of estimated coefficients, given the short spans of our data set. For constructing an individual country-group model, we assume that most economies, with the possible exception of the USA, are small relative to the world economy so that foreign variables satisfy the property of weak exogeneity. In the empirical analysis that follows, we use the Akaike information criterion (AIC) with a maximum of 2 lags to choose the lag order of the domestic variables, and we set the lag order of the foreign variables equal to one. The short lags are due to data limitations. Based on these criteria, we choose an augmented vector autoregressive (VARX*) model, with a second-order dynamic specification of domestic variables and a first-order dynamic specification of country-specific foreign variables, for all country-group models. The resulting VARX*(2,1) specification for each country group can be written as

$$g_{i,t} = \Phi_1 g_{i,t-1} + \Phi_2 g_{i,t-2} + \Psi_0 g^*_{i,t} + \Psi_1 g^*_{i,t-1} + u_{i,t}, \tag{28}$$

for $t = 1, 2, ..., T$, and $i = 1, 2, ..., N_j$ where N_j is the number of countries in group j. The notation is such that $g_{i,t}$ is a $k_i \times 1$ vector of endogenous/domestic variables, $g^*_{i,t}$ is a $k^*_i \times 1$ vector of exogenous/country-specific foreign variables, Φ_1 and Φ_2 are $k_i \times k_i$ matrices of coefficients associated with lagged endogenous variables, Ψ_0 and Ψ_1 are $k_i \times k^*_i$ matrices of coefficients associated with country-specific foreign variables, and $u_{i,t}$ is a $k_i \times 1$ vector of reduced-form residuals with a variance-covariance matrix $\Sigma = E[u_{i,t}u'_{i,t}]$ for all i and all t. Country-specific fixed effects are allowed in our model via the inclusion of country-specific dummy variables. However, we have dropped the intercept and country-specific dummy variables from equation (28) to simplify the notation. Note that we do not include the USA in any country group when we estimate the model, because it is likely that the weak exogeneity assumption is violated, given the size of the US economy and the importance of the USA to global economic interaction.

The corresponding conditional vector error correction model (VECM) is given by

$$\Delta g_{i,t} = -\Pi z_{i,t-1} - \Phi_2 \Delta g_{i,t-1} + \Psi_0 \Delta g^*_{i,t} + u_{i,t}, \tag{29}$$

where

$$\Pi = (I - \Phi_1 - \Phi_2, -\Psi_0 - \Psi_1) \text{ and } z_{i,t-1} = \left(g'_{i,t-1}, g^{*\prime}_{i,t-1}\right)'.$$

It is easy to see that the cointegrating relation among variables is summarized in a $k_i \times (k_i + k^*_i)$ matrix Π. Suppose that the rank of Π is $r_i \leq k_i$, implying that there are r_i long-run relationships among the variables.

The matrix $\Pi = \alpha\acute{\beta}$, where α is a $k_i \times r_i$ loading matrix of full column rank and β is a $(k_i + k_i^*) \times r_i$ matrix of cointegrating vectors of rank r_i. For the estimation of the vector error correction model (VECM), we will impose cointegrating restrictions $\widehat{\beta}$ computed by using panel dynamic OLS estimation, or $\widehat{\beta}$ suggested by economic theory. Thus we can rewrite equation (29) as

$$\Delta g_{it} = B_0 ecm_{i,t-1} + B_1 \Delta g_{i,t-1} + B_2 \Delta g^*_{i,t} + u_{i,t} \tag{30}$$

where $ecm_{i,t-1} = \acute{\widehat{\beta}} z_{i,t-1}$, $B_0 = -\alpha$, $B_1 = -\Phi_2$, $B_2 = \Psi_0$.

For estimation, we apply within-group estimation i.e. we include a constant and country-specific dummy variables in the VECM to account for country-specific fixed effects. Alvarez and Arellano (2003) show that when T/N tends to a positive constant, the within-group estimator has negative asymptotic biases of order $1/T$ and these biases disappear when $N/T \to 0$, while the GMM estimator is asymptotically biased of order $1/N$. They also find that the crude GMM estimator which neglects autocorrelation in the error terms is inconsistent, despite being consistent for fixed T. Moreover, Judson and Owen (1999) show that based on a root mean squared error criterion, the within-group estimation performs better than others when $T = 30$ and $N = 20$. According to these findings, the within-group estimation seems to be more appropriate than other estimation techniques for our data set for which $T = 39$ and $N = 8$ on average for each country-group model, because the downward bias is relatively insignificant.

As mentioned before, the weak exogeneity condition of country-specific foreign variables is the main assumption underlying the estimation strategy of the model for each country group. Following the notions of long-run causality and long-run forcing, suggested by Johansen (1992) and Granger and Lin (1995), the weak exogeneity condition means that there is no long-run feedback from a set of domestic variables g_{it}, to foreign variables g^*_{it}, regardless of lagged short-run effects between these variables. We can indirectly test this condition by taking the marginal model of the country-specific foreign variables that include estimated error-correction terms from DOLS, and then using a standard F test to test the joint significance of the estimated error-correction terms. To illustrate, the weak exogeneity test of the e^{th} element of the foreign variable vector $g^*_{e,i,t}$, in the model of each country group can be performed by estimating

$$\Delta g^*_{e,i,t} = c_e + \sum_{n=2}^{N_j} \lambda_{e,n} d_n + \sum_{R=1}^{r_j} \phi_{e,R} ecm^R_{i,t-1} + \sum_{p=1}^{P_j} \theta_{e,p} \Delta g_{i,t-p} + \sum_{m=1}^{M_j} \sigma_{e,m} \Delta g^*_{i,t-m} + \xi_{e,i,t}, \tag{31}$$

for $t = 1, 2, ..., T$, and $i = 1, 2, ...N_j$ where N_j is the number of countries in group j, d_n are country-specific dummy variables, c_e is a constant term and $\xi_{e,i,t}$ is an error term. Note that $ecm^R_{i,t-1}$, $R = 1, ..., r_j$ are estimated error-correction terms associated with r_j cointegrating vectors estimated by DOLS for the j^{th}

country group. For testing the weak exogeneity of e^{th} element of the foreign variable vector, we test the null hypothesis that $\phi_{e,R} = 0$, $R = 1, ..., r_j$ using a standard F test. In our study, the lag orders are chosen in the light of serial correlation tests and available observations. We allow the weak exogeneity tests to include a maximum of lags 5 for both domestic and exogenous variables.

Given the VARX* model in (28), our interest is to examine impulse responses to economically meaningful structural shocks $v_{i,t}$. This leads to the problem of identification, given the correlation among reduced-form residuals $u_{i,t}$. The identification issue is how to decompose reduced-form residuals $u_{i,t}$ into structural shocks $v_{i,t}$. Recall that the dimension of $g_{i,t}$ is k_i. We adopt an appealing assumption in the VAR literature that there are k_i fundamental innovations $v_{i,t}$, which are mutually independent and normalized to be of variance 1, i.e. $E(v_{i,t}v_{i,t}^{'}) = I_{k_i}$. To obtain independence of the fundamental innovations, we find a matrix A such that $u_{i,t} = Av_{i,t}$. We can then rewrite the reduced-form VARX* into the structural VARX* as below:

$$A^{-1}g_{i,t} = A^{-1}\Phi_1 g_{i,t-1} + A^{-1}\Phi_2 g_{i,t-2} + A^{-1}\Psi_0 g_{i,t}^{*} + A^{-1}\Psi_1 g_{i,t-1}^{*} + v_{i,t}. \tag{32}$$

It is clear that if we want to estimate the impulse responses of endogenous variables to structural shocks $v_{i,t}$, we require the identifying matrix A. Note that the immediate impact or impulse vector of the j^{th} structural innovation (which is the j^{th} element of the vector $v_{i,t}$), of a one standard deviation shock in $v_{i,t}$ on each endogenous variable in the system can be represented by the j^{th} column of the matrix A, a_j.

The current set of restrictions that we have for specifying the matrix A is $\Sigma = E[u_{i,t}u_{i,t}^{'}] = AE[v_{i,t}v_{i,t}^{'}]A^{'} = AA^{'}$. Such restrictions are not sufficient to achieve a unique solution for the matrix A. In particular, we need at least $\frac{k_i(k_i-1)}{2}$ additional identifying restrictions to be imposed on the matrix A.

There are three identification procedures commonly used for the orthogonalization of shocks. One procedure defines the matrix A as a lower triangular Cholesky factor of Σ. This matrix depends on a recursive ordering of endogenous variables in the system. The second procedure imposes some structural relationships between structural shocks $v_{i,t}$ and reduced-form disturbances $u_{i,t}$ that are implied by some theoretical models or economic intuition. This can be achieved by imposing short-run and long-run restrictions such as restrictions on temporary and permanent components[7]. The third procedure is a novel approach and is used in our study. This procedure identifies a shock by imposing sign restrictions on some of the impulse response functions.

[7] See Blanchard and Quah (1989) for more details.

3.2 Sign Restrictions with a Penalty-Function Approach

Traditional identifications are commonly based on zero contemporaneous and/or long-run restrictions. The results from these approaches depend on the chosen decomposition of a variance-covariance matrix, i.e. the restrictions used for reordering variables, or used for selecting a particular Cholesky decomposition can alter the responses to a shock. This leads to the imposition of very stringent restrictions, most of which do not rely on theoretical considerations.

In this paper, we avoid these problems by imposing sign restrictions on the impulse response functions to identify four types of underlying disturbances; productivity growth, trade liberalization, monetary policy, and government spending shocks. This new identification strategy was developed by Uhlig (2005) and extended by Mountford and Uhlig (2009). With this approach, it is not necessary to impose a priori zero contemporaneous impacts or long-run restrictions of shocks. Instead, it requires only a set of economically plausible restrictions that are often used implicitly by researchers. It makes a priori theoretical restrictions explicit and leaves the question of interest open. Underlying shocks can be identified by examining whether the signs of the corresponding impulse responses are accepted by a priori censensus considerations. With this approach, we also obtain results that are robust to reordering variables.

As shown by Uhlig (2005) and generalised by Mountford and Uhlig (2009), the identification does not depend on any particular matrix A. On the one hand, if there exists a k-dimensional vector m of unit length such that $a_j = \overrightarrow{A}m$, where $\overrightarrow{A}\overrightarrow{A}' = \Sigma$ and $\overrightarrow{A}$ is any arbitrary decomposition of Σ such as a lower triangular Cholesky factor, we can obtain an impulse vector a_j even though the true matrix A is not identified. In general, we can write $A = [a^{(1)}, ..., a^{(k_i)}] = \overrightarrow{A}M$ where $M = [m^{(1)}, ...m^{(k_i)}]$ is an orthonormal matrix such that $MM^{'} = I_{k_i}$. Uhlig (2005) uses this property to show that the impulse response $r_a(h)$ at horizon h to the impulse vector a can be computed as a linear combination of the impulse responses obtained using the Cholesky decomposition of Σ. This can be represented as

$$r_a(h) = \sum_{j=1}^{k_i} m_j r_j^c(h),$$

where $r_j^c(h) \in \mathbb{R}^{k_i}$ is a $k \times 1$ vector of the impulse response at horizon h to the j^{th} shock in a Cholesky decomposition of Σ, i.e. the j^{th} column of $\overrightarrow{A}$.

Moreover, it is interesting to perform variance decompositions or in other words, compute how much the shock contributes to the variance of the h-step ahead forecast error. The fraction $\phi_{a,c,h}$ of the variance of the h-step ahead forecast revision $E_t[g_{t+h}] - E_{t-1}[g_{t+h}]$ for variable c, explained by shocks corresponding to the impulse vector a can be obtained by

$$\phi_{a,c,h} = \frac{(r_{a,c}(h))^2}{\sum_{j=1}^{k_i}(r_{j,c}(h))^2},$$

where $r_{j,c}(h)$ denotes the impulse response of variable c at horizon h to the j^{th} shock.

The sign restriction strategy can be applied to achieve the identification of s structural shocks, where $s \leq k_i$. As our study focuses on four underlying shocks, we need to characterize an impulse matrix $[a^{(1)}, a^{(2)}, a^{(3)}, a^{(4)}]$ of rank 4, rather than all impulse vectors in the matrix A. In particular, we draw a $k_i \times 4$ matrix $M = [m^{(1)}, m^{(2)}, m^{(3)}, m^{(4)}]$ which contains four orthonormal vectors, and then we can calculate a $k_i \times 4$ impulse matrix as $[a^{(1)}, a^{(2)}, a^{(3)}, a^{(4)}] = \overrightarrow{A}M$. By construction, the covariances between any pair of the underlying shocks $v_t^{(1)}, v_t^{(2)}, v_t^{(3)}$ and $v_t^{(4)}$ corresponding to the impulse vectors $a^{(1)}, a^{(2)}, a^{(3)}$ and $a^{(4)}$ are zero, so we can characterize such an impulse matrix by imposing economically meaningful sign restrictions on the impulse responses.

Following Uhlig (2005) and Mountford and Uhlig (2009), we use a Bayesian method to implement the sign restrictions. This is because it simplifies computation of the responses and their error bands. The posterior is given by the usual Normal-Wishart posterior for the VECM parameters (B, Σ). A Monte Carlo integration is performed. Given the estimated vector error correction model, we take a joint draw from the posterior of the Normal-Wishart distribution for (B, Σ) and a draw from a uniform distribution over the unit sphere for candidate m vectors. The Cholesky decomposition factor, $\overrightarrow{A}$ is computed using a draw of Σ from the posterior. Consequently, we can calculate the candidate impulse responses as $a_j = \overrightarrow{A}m$.

In this paper, we follow a sign restriction approach with a penalty function, rather than a pure-sign restriction approach. The main difference between these two approaches is that with a pure-sign restriction approach, all impulse vectors satisfying the sign restrictions are considered equally for determining the impulse responses, while the penalty-function approach chooses the best of all impulse vectors via the use of a criterion function. Thus, with the penalty-function approach, it is possible to obtain impulse response functions with smaller standard errors.

In particular, for a pure-sign restriction approach, if all signs of the candidate impulse response correspond to all sign restrictions imposed on each structural shock, we keep the draw. Otherwise, such a draw is discarded. However, for the penalty function approach, we find the best impulse vector a for any given (B, Σ). Although no impulse response might satisfy all sign restrictions, the impulse vector a which generates responses that satisfy the sign restrictions as closely as possible is considered. This can be achieved by minimizing a criterion function. We repeat the draws until we obtain enough impulse responses and then find the median as well as the 16 percent and 84 percent quantiles for the sample of impulse responses.

Note that these two quantiles give one standard deviation error bands, given that variables follow a normal distribution.

The penalty function suggested by Uhlig (2005) is

$$f(w) = \begin{cases} w & \text{if} \quad w \leqslant 0, \\ 100 \times w & \text{if} \quad w \geqslant 0. \end{cases} \tag{33}$$

This function is asymmetric when imposing sign restrictions i.e. we penalize positive/wrong responses 100 times more than we reward negative/correct responses. We minimize the criterion function in order to find the best impulse vector a for each draw of (B, Σ) from the posterior. Thus

$$a = \arg \min_{a=\overrightarrow{A}m} \Psi(a)$$

and

$$\Psi(a) = \sum_{l} \sum_{h=0}^{H} f(\iota_l \frac{r_{l,a}(h)}{\sigma_l})$$

where l is the set of variables that are sign-restricted for the identification of shocks, $r_{l,a}(h)$ is the impulse response of the l^{th} variable at horizon h to an impulse vector a, and σ_l is the standard deviation of the first-differenced variable l. Note that σ is used for rescaling impulse responses, or in other words, generating standardised impulse responses so that the deviations across different impulse responses are comparable to each other. Define $\iota = -1$ if l belongs to the set of variables that respond positively to a given shock according to a sign restriction, and $\iota = 1$ if l belongs to the set of variables that respond negatively to a given shock. Using numerical minimization, we can identify the first shock $a^{(1)} = \overrightarrow{A}m^{(1)}$ and use this to calculate the impulse responses.

The identification of the second shock can then be achieved by

$$a = \arg \min_{a=\overrightarrow{A}m, m'm^{(1)}=0} \Psi(a).$$

We can identify an impulse matrix $[a^{(1)}, a^{(2)}]$ by adding the restriction that the second shock is orthogonal to the first shock. After we identify the first two shocks, i.e. $a^{(1)} = \overrightarrow{A}m^{(1)}$ and $a^{(2)} = \overrightarrow{A}m^{(2)}$, we can identify the third shock using

$$a = \arg \min_{a=\overrightarrow{A}m, m'm^{(1)}=0, m'm^{(2)}=0} \Psi(a).$$

This ensures that the minimisation chooses an a that is orthogonal to the first two shocks. Likewise, additionally restricting orthogonality to the first three shocks, the fourth shock is identified by minimizing the problem below:

$$a = \arg \min_{a=\overrightarrow{A}m, m'm^{(1)}=0, m'm^{(2)}=0, m'm^{(3)}=0} \Psi(a).$$

Following Uhlig (2005), we use a Bayesian approach to perform the computation. That is, for each draw from the posterior of (B, Σ), we can identify the shocks using the criterion function above. We take 1,000 draws from the posterior and then we compute the median impulse responses as well as one standard deviation error bands. These can be obtained from the impulse responses at the 16, 50, and 84 percent quantiles, given that the variables follow a normal distribution.

3.3 Identifying Assumptions and Implementation based on Sign Restrictions

We adopt all sign restrictions displayed in Table 1. All restrictions are imposed as $\leq (-)$ or $\geq (+)$ and they are in line with the theoretical literature and standard international macroeconomic theory. Also, they have been commonly used in the empirical literature that uses the sign restriction approach.

Table 1: Identifying Sign Restrictions

Shocks	q	x	y	gov	$open$	si	π
Productivity Improvement Shock		+	+				
Trade Liberalization Shock					+		
Contractionary Monetary Policy Shock						+	−
Government Spending Shock				+			

In this table, q represents the real effective exchange rate, x is the traded-nontraded productivity differential, y is real GDP, gov is the government consumption share, $open$ is the degree of openness in the economy, si is the nominal short-term interest rate, and π is the inflation rate.

The productivity improvement shock can be viewed as a supply shock, because it is well known in the macroeconomic literature that a positive shock to productivity leads to a significant change in real variables. In our study, the productivity shock is identified as a shock that causes productivity in traded sectors relative to nontraded sectors and real output to increase for a year. This corresponds to the economic literature such as the BSH, and a general equilibrium model of an open economy such as the 1-2-3 model, developed by Devarajan, Lewis, and Robinson (1990). We do not impose any restrictions on other variables as there is no macroeconomic theory explaining the exact responses of these variables to the productivity shock.

We identify the trade-liberalization shock as an unexpected rise in the ratio of the sum of imports and exports to GDP, denoted by the variable called open, for a year, and it is orthogonal to the productivity shock, with no sign restrictions placed on other variables.

We have constructed the contractionary monetary policy shock so that it is orthogonal to the productivity shock and the trade liberalization shock. We also consider this shock as a consequence of a surprise increase in the nominal interest rate and a contemporaneous fall in inflation over a year. Although these conditions can be controversial, i.e. a presence of the price puzzle is raised by Sims (1992), the sign restrictions imposed here are the same as those in the empirical works by Uhlig (2005) and Peersman (2005).

The last shock is the government spending shock. This shock is identified by restricting the impulse responses of the government consumption variable to be positive for a year after the shock, while the responses of other variables to the shock are not restricted at all. In the literature, some researcheres use this variable to represent a fiscal policy. This may not be actually true as the fiscal policy does not only depend on government spending but also on government revenue. Similar to Mountford and Uhlig (2009), we control for the productivity shock (or a supply shock), the monetary policy shock, and the trade liberalization shock when identifying this shock. In particular, we have constructed the government spending shock so that it is orthogonal to the productivity improvement shock, the trade liberalization shock, and the contractionary monetary policy shock.

In short, we firstly identify a productivity shock and then identify a trade liberalization shock which is orthogonal to the productivity shock via the imposition of sign restrictions. Next the monetary policy shock is identified via sign restrictions and by imposing orthogonality to the productivity shock and the trade liberalization shock. Finally, we identify a government spending shock as a shock that is orthogonal to the productivity shock, the trade liberalization shock, and the monetary policy shock. The idea behind this ordering is that it is difficult to distinguish the movement of each variable caused by a shock to that variable from the contemporaneous movement in that variable caused by other shocks. The orthogonality condition can help to identify the shock by filtering out the contemporaneous responses of each variable to other shocks. In our study, we decide to begin with the shock to productivity as it seems implausible that other shocks can affect productivity in the short run. Then we choose the trade liberalization shock as a second shock for identification. This is because the open variable can be viewed as a trade-policy variable which takes time to adjust, similar to the productivity variable. Nonetheless we have checked that our results are robust to change in the order of these first two shocks. For monetary policy and government spending shocks, we choose to order these shocks after the productivity shock and the trade-liberalization shock in order to filter out the effects of business cycle and trade liberalization shocks. Also this ordering corresponds to that of Mountford and Uhlig (2009).

We have attempted to be consistent with theoretical literature and empirical research in choosing our sign restrictions. We do not impose any responses of uncertain sign. As our focus is on the responses of the real exchange rate to a shock to exchange rate determinants and policy applications, we leave all its responses unrestricted.

As mentioned in the previous section, we use a penalty function approach to apply sign restrictions. In addition to the above sign restrictions, we impose a one period constraint as well, i.e. the impulse responses are required to satisfy the sign restrictions both contemporaneously and over the next year. However, we find that using different horizons over which the sign restrictions are binding does not have much influence on the results.

For computation, we find a "best impulse matrix" by undertaking the numerical minimization of the above criterion function $\Psi(a)$ on the unit sphere, given each draw of B and Σ from the posterior. We parameterize the space of unit-length vectors by using the stereo projection[8], which is available in the RATS statistical package to implement this minimization. We do the minimization procedure twice for each draw, starting it from two different initial random vectors in order to check whether the best impulse vector we obtained is the optimal solution. In particular, we examine whether the two minima found are very close or the same. If they are the same or different by less than 0.01, we keep the impulse vector. In contrast, if they generate values of the total penalty that differ by more than 0.01, we keep only the vector which generates the smaller value of the total penalty, and we discard the other. Therefore given each draw of B and Σ, we will obtain a selected impulse matrix for computing impulse responses. Then we draw a new B and Σ, and start a new minimization procedure using the last set of minimizers as one of initial vectors. We continue and repeat these procedures until we have acquired 1,000 draws of B and Σ generating 1,000 best impulse matrices. This generates a sample of 1,000 impulse responses that satisfy the sign restrictions. Given this sample, we find the impulse responses at the 16, 50, and 84 percent quantiles for each of 9 step-ahead forecasts. This gives the median impulse responses as well as one standard deviation error bands, assuming that variables follow a normal distribution.

4 Empirical Investigation

In this section, we illustrate our methodology and show the results. Most of the analysis is based on a vector error-correction model with exogenous variables, and we adapt it to a set of country groups in order to take advantage of panel techniques. This reduces the limitations associated with small sample size, which is an important problem when studying developing countries, especially the Asian countries that are of particular

[8] The stereo projection is a way of drawing the unit sphere onto the plane through the equator. With this technique, the angles at which curve cross each other are preserved but areas and distances are distorted.

interest in this thesis.

4.1 Data Description

We employ an unbalanced panel data set that includes annual time series from 1970-2008. The data set covers 30 countries. Five countries that joined the European monetary union in 1999 are included in our country set, and they are grouped together into a single Euro area economy when estimating the model. We employ the data set used to study the BSH in the first chapter, but add in some new series of interest. However, in this chapter we drop Myanmar, Brunei Darussalam, and Vietnam from our country set because their data sets miss several observations. Our country grouping also differs from before because in addition to the classification using the level of economic development, each group of countries, developed and developing, is separated by regions. In particular, we classify our country set into four groups: European countries, other developed countries, Asian developing countries, and non-Asian developing countries. The reasons for this are to group similar countries together and to achieve credible results from applying panel estimation to each country group. See Appendix A.1 for details of the countries in each group.

Throughout our analysis, the core variables considered are the log real effective exchange rate[9] ($q_{it} = (e_{it} - p_{it}) - (e^*_{it} - p^*_{it})$ where e_{it} is the log nominal exchange rate with respect to US and p_{it} is the log general price index, measured by CPI, for country i during the period t), and seven main real exchange rate fundamentals associated with the BSH framework and the Edwards model i.e. the log traded-nontraded productivity differential (x_{it}); the log relative price of nontraded goods to traded goods (rp_{it})[10]; log real GDP (y_{it})[11]; the log terms of trade (tt_{it}), defined as the ratio of the export price index to the import price index; the log government consumption share (gov_{it}), measured as the ratio of government consumption to GDP; the log openness of the economy ($open_{it}$), measured as the ratio of the sum of exports and imports to GDP; the real interest rate (r_{it}), defined as the nominal interest rate ($si_{it} = \ln(1 + NI_{it}/100)$) minus the inflation rate ($\pi_{it} = p_{it} - p_{i,t-1}$) where NI_{it} is the short-term nominal interest rate per annum measured as a percentage and p_{it} is the log general price index, measured by CPI[12]. In addition to these real exchange rate determinants, we include the nominal short-term interest rate (si_{it}) to investigate the responses of real exchange rates to a monetary policy shock, and include the log of oil prices (oil_t) to account for global

[9] From our measurement of the real effective exchange rate, an increase in the real exchange rate represents a depreciation.

[10] We follow the industry classification approach used in the first chapter for constructing the productivity and the relative price of nontraded goods to traded goods. The price index for each sector is defined as the value added in current local currency units divided by the value added in constant local currency units.

[11] We include log real GDP rather than per capita GDP used in the first chapter because we want to investigate the relationship between the real exchange rate and economic growth in this study, while the focus of the previous study is to examine whether per capita GDP widely used in the BSH literature is a good proxy for the traded-nontraded productivity differential.

[12] We follow Pesaran et al. (2004) for constructing the real interest rate variable.

unobserved factors. Also another variable included is the inflation rate (π_{it}), as this variable may influence real exchange rates in the short run. Data sources are provided in Appendix A.2.

We group five European countries together into a single euro area regional economy, since they have similar economies. The regional variables of this group are constructed from country-specific variables using weighted averages based on average PPP-adjusted GDP series over the period 2002-2008[13]. We construct the region-specific foreign variables for this region in the usual way, as mentioned in Section 3.

4.2 Unit Root Tests

We start with examining the integration properties of the series to determine appropriate transformations of the variables for use in the country/region-specific error correction models. We conduct the Augmented Dickey-Fuller (1979) tests with generalized least squares detrending (ADF-GLS), developed by Elliott et al. (1996) for the levels and first differences of the domestic and country/region-specific foreign variables, as well as of the differential variables between domestic and foreign variables on each cross-sectional unit[14]. This test improves small-sample power relative to standard ADF tests.

Given the limitations of our short data set, we also perform panel data unit-root tests, which provide dramatic improvement in statistical power by using cross-sectional information. The cross-sectional augmented Dickey-Fuller (CADF) version of the t-bar test proposed by Im, Pesaran and Shin (IPS) (2003), the inverse normal combination test suggested by Choi (2001), and the inverse chi-squared (Fisher) test proposed by Maddala and Wu (1999), together with GLS detrending developed by Elliott et al. (1996) are employed to achieve satisfactory size and power for small values of N and T. These tests investigate the null hypothesis that all cross-sectional series in the panel are unit root processes against the alternative hypothesis that at least a significant fraction of the individual series in the panel is stationary. The null for these tests, except the inverse chi-squared test, will be rejected with large enough negative values of its test statistics. Note that the lag length p for the tests is selected by the Bayes information criterion (BIC), allowing a maximum of 4 lags. We choose to use a truncated version of the individual CADF statistics and their p-values to avoid the impacts of extreme values. Then the cross-sectionally augmented IPS (CIPS) test value is a simple average of the individual CADF statistics. The inverse chi-squared and the inverse normal tests are constructed as is common practice in meta-analysis: combining p-values from a test of each cross-sectional unit in the panel.

Tables B2-B4 in Appendix B report the test results for both time-series and panel unit root tests. The

[13] The PPP-GDP series are from WDI data, World Bank. Moreover, as nominal exchange rates in these five countries are the same after 1999, in order to obtain an appropriate euro area exchange rate over the period 1970-2008, we use a weighted average of monthly exchange rate series, computed by Anderson et al (2011).

[14] We also apply ADF unit root tests to all series, and find similar results. The results of these tests are available from the author upon request.

test results are in line with what is known in the literature. They show that most variables, i.e. the real exchange rate, traded-nontraded productivity, the relative price of nontraded goods, terms of trade, real GDP, government consumption, nominal interest rates, oil price and openness, in levels and differentials (the differences between domestic and country/region-specific foreign variables) are approximately I(1).

However, the individual and panel test statistics seem to suggest that real interest rate differentials ($r_{it}^* - r_{it}$) are stationary. This implies that there is no cointegrating relationship between real interest rate differentials and real exchange rates. As we focus on long-run real exchange rate movements, the real interest rate differentials (and real interest rates) are not considered further, and are dropped from our model. Moreover, the test results for inflation variables are ambiguous: Foreign-inflation variables are likely to be I(1) according to both time-series and panel unit root tests, while the evidence on domestic-inflation variables is less clear cut. The domestic-inflation variable is possibly I(0) in some countries, especially in Asia. Pesaran et al. (2004) include inflation instead of price in their models, because including I(2) series in the model can lead to more serious problems than using overdifferenced series. We do the same here. In short, we treat all of these domestic and foreign variables as well as the differentials between domestic and foreign variables as I(1), with the exception of real interest rate differentials in all countries and domestic-inflation variables in Asian developing group, which we regard as I(0).

4.3 Testing for the long-run relationships

In this section, we investigate the existence of long-run cointegrating relationships between the variables, and focus on the issue of identifying these long-run relationships. Our identification strategy relies on economic theory, focusing on the long-run relationships between variables in equilibrium. In the context of real exchange rates, the relationship is derived from the BSH and from inter-temporal optimization conditions in the general equilibrium framework of the Edwards model discussed above. Based on the long-run properties of macroeconomic models, we consider the following long-run relationships as possible candidates:

Purchasing Power Parity	q_{it}	$\sim$	$I(0)$	(1)
Fisher Equation	$si_{it} - \pi_{it}$	$\sim$	$I(0)$	(2)
Output Convergence	$y_{it} - y_{it}^*$	$\sim$	$I(0)$	(3)
Uncovered Interest Parity	$si_{it} - \sigma si_{it}^* - E(\Delta e_{i,t+1}^*)$	$\sim$	$I(0)$	(4)

Combination of the Balassa-Samuelson and Edwards models

$$q_{it}-\lambda_{i1}(x^*_{it}-x_{it})-\lambda_{i2}(rp^*_{it}-rp_{it})-\lambda_{i3}(gov^*_{it}-gov_{it})-\lambda_{i4}(y^*_{it}-y_{it})-\lambda_{i5}(r^*_{it}-r_{it})-\lambda_{i6}(tt_{it})-\lambda_{i7}(open_{it}) \sim I(0) \quad (5)$$

The first relationship, called PPP is the well-known theory of long-term equilibrium exchange rates based on the relative price levels between countries. The second relationship is the Fisher Equation which shows the relationship between nominal, real interest rates and inflation. It was first proposed by Irving Fisher (1977) in his work "the theory of interest rate". The third relationship represents the relative output convergence condition derived from the Solow-Swan neoclassical growth model. The fourth relationship is the modified version of uncovered interest parity (UIP) condition which relates nominal interest rates and market expectations of future exchange rates. Since the results of the unit root tests in Table B3, Appendix B, show that $E(\Delta e^*_{i,t+1})$ is $I(0)$, this relationship can be reduced to $si_{it} - \sigma si^*_{it} \sim I(0)$. In this paper, we focus on the fifth relationship. It shows the relationship between the real exchange rate and its main fundamentals. As the movement of real exchange rates depends not only on domestic impacts but also on external impacts from outside countries, we choose to use variables in relative terms. However, we do not use relative terms for the terms of trade and openness variables because they have already accounted for the interaction between domestic and foreign countries by construction.

We first investigate the existence of cointegrating relationships between variables in our model. In the literature, the common way to deal with this is to apply the system approach, i.e. the traditional Johansen (1988) cointegration test, in order to determine the number of cointegrating vectors in each individual VARX* model. However, with a large set of variables and a small sample in our data set, this method is infeasible. More parsimonious approaches are considered for this reason. We turn to single-equation cointegration tests inspired by the above theoretical long-run relationships, so as to avoid small-sample size problems associated with system-based tests. We perform several tests based on single-equation frameworks together with critical values computed from empirical distributions in order to test the above theoretical relationships[15]. The results show that there is little evidence of long-run relationships between the variables, inconsistent with economic theory. In order to ascertain whether this is a consequence of low power when applying the test to series with short length, we conduct the KPSS test for the null of stationarity proposed by Kwiatkowski et al. (1992) and Shin (1994) and find that the null hypothesis cannot be rejected for most countries, implying the presence of cointegration. This contradiction indicates that although we avoid the problems associated with the system approach, the test results from single-equation cointegration tests are

[15] The Engle and Granger (1987) cointegration test, the weighted symmetric cointegration test introduced by Cook and Vougas (2008), the error-correction test based on the lagged dependent variable by Banerjee (1998), and the bounds testing approach developed by Pesaran et al. (2001) are used.

still unreliable.

Panel techniques are applied in order to exploit more available information from cross-sectional units in the sample. We can implicitly examine the relationship (1) by conducting panel unit root tests on the real exchange rate (q_{it}). The results, shown in Table B2, Appendix B, show that the PPP condition does not hold in any country, corresponding to the explanation in the literature that other real exchange rate determinants can lead to persistent deviations from long-run equilibrium[16].

In order to test for the presence of long-run relationships (2)-(5) in a panel of countries, we apply seven residual-based tests for the null of no cointegration in dynamic heterogeneous panels, suggested by Pedroni (1999). These panel cointegration tests are based on the null hypothesis that for each member of the panel the variables of interest are not cointegrated, while the alternative hypothesis is that there exists a single cointegrating vector for each country in panel. With these approaches, the cointegrating vectors are allowed to be different for each country. Since the tests are based on the assumption of cross-sectional independence in the error term, we include a set of common time dummies in the hypothesized cointegrating regression to accommodate some forms of cross-sectional dependence across the different members of the panel so that the remaining idiosyncratic error terms become independent.

Our test results show that the seven test statistics lead to different conclusions. According to Pedroni (1999), the comparative advantage of each test statistic in terms of small sample size and power properties will depend on the underlying data-generating process (DGP). Pedroni (2004) shows that in a situation similar to ours, i.e. for $N = 20$ cross-sectional units and $T = 20 - 40$ time periods, panel-t statistics and group-t statistics have higher power than other test statistics, given a DGP of $\phi = 0.90 - 0.95$ where ϕ is the autoregressive coefficient of residuals. According to this finding, we will focus on the results of the parametric panel t-statistics, the nonparametric panel t-statistics, the parametric group t-statistics, and the nonparametric group t-statistics.

The results of the Pedroni (1999) cointegrating tests for the Fisher Equation, the output convergence relation, and the UIP condition are shown in Tables B5.1-B5.3, Appendix B. We find that there is a cointegrating relationship between the nominal interest rate and inflation, corresponding to the Fisher Equation, in developing countries[17], whereas this relationship is not found in developed countries. Also, the results show that the output convergence does not hold in all country groups. This is in line with the results of unit root tests on $(y^*_{it} - y_{it})$ displayed in Table B4, Appendix B. Interestingly, we find evidence for the UIP condition

[16] See, for example, Pesaran (2007), Pesaran et al. (2009), and Bussière et al. (2009).

[17] However, according to the results of unit root tests in Table B2, Appendix B, we treat the inflation as a stationary variable in the model of developing-Asia group.

in most country groups. This finding corresponds to what we expect, as nowadays financial markets in most countries are integrated with each other, and have become more like a global financial market.

The real interest rate differential ($r^*_{it} - r_{it}$) is dropped from the test of relationship (5) as the unit root tests indicate that it is stationary. However, (5) still relates to a large set of variables (eleven variables) and the span of the data sets in each country group is short. Given these potential problems, we perform the Pedroni (1999)'s cointegration tests for (5) using the whole country set in order to increase the power of the test. This is different from testing relationships (2)-(4), as in these relationships we conduct the tests separately for each country group. Table 2 reports the results of the panel and group-mean statistics for the null of no cointegration in case of the long-run relationship (5). All tests support the existence of a cointegrating relationship between the real exchange rate and its fundamentals.

Table 2: Cointegration tests for a relationship between the real exchange rate and its 6 fundamentals

Variables	Panel PP	Group PP	Panel ADF	Group ADF
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}),$ $(y^*_{it} - y_{it}), tt_{it}, open_{it}$	0.00	0.00	0.00	0.03

Note: 1) The number in the table stands for the P-value for the tests under the null hypothesis of no cointegration.

2) The lag length in ADF-type regression of Pedroni's tests is selected by BIC with a maximum of 4 lags.

3) The critical value is determined by the standard normal distribution.

Next we examine whether the collection of variables in this cointegrating relationship contains redundant elements. In addition to the issue of parsimony, a cointegrating vector including an irrelevant variable can lead to an inconsistent estimate of the cointegrating relationship. Davidson (1998) discusses this problem, and accordingly defines an irreducible cointegrating relation as a set of cointegrated variables, none of which can be omitted without loss of the cointegration property.

Davidson (1998) introduces a technique in order to obtain the irreducible cointegrating relation. As we start with a test of cointegration in a general model, the method we use to obtain the irreducible cointegrating relation is called the "general-to-particular" mode of search. For illustration, we begin with a large set of cointegrated variables comprising the real exchange rate and six fundamentals as shown in Table 2. Then we drop one fundamental at a time and conduct a Pedroni cointegration test on the remaining variables. If the remaining variables are still cointegrated after dropping one variable, this implies that the omitted variable is irrelevant and it can be omitted. In contrast, if the remaining variables are not cointegrated, this indicates that the omitted variable is important for the cointegrating relation and cannot be dropped. We repeat this for every possible subset of remaining variables until we find the collection of variables where no variables can be omitted without loss of the cointegration property.

In our study, we find that the relative price of nontraded goods to traded goods and the terms of trade can be omitted without loss of cointegration. In other words, the relative price and the terms of trade can be viewed as "irrelevant" variables and can be eliminated from the cointegrating relationship. The results of these tests are displayed in Tables B6.1-B6.3, Appendix B. From this, the remaining set of variables i.e. the real exchange rate, the productivity gain in traded sector, real GDP growth, the openness of the economy, and government consumption can be regarded as the irreducible cointegrating relation in a panel of countries.

However, a panel cointegration test above can provide only guidance regarding the existence of a cointegrating relationship among the variables of interest in the whole data set. In order to understand the feature of this cointegrating relationship in more details, we then separately test whether the irreducible cointegrating relation holds in each country group. Table 3 shows these test results for each country group.

Table 3: Cointegration tests for the real exchange rate relationship in each country group

Country groups	Panel ADF	Group ADF
European countries	0.05	0.05
Other developed countries	0.06	0.04
Asian developing countries	0.52	0.58
Non-Asian developing countries	0.00	0.00

See Table 2.

With the smaller sets of countries, we focus only on panel ADF and group ADF statistics as Pedroni (2004) shows that they are the most powerful for very small N and T. As shown in Table 3, it is clear that the cointegrating relationship between the real exchange rate and its four fundamentals is still satisfied in European, other developed-country, and non-Asian developing-country groups. In contrast, there is not enough evidence in favor of this cointegrating relationship in the Asian developing-country group. However, it is important to note that as the tests are based on $N = 8$, the panel test statistics might have low power and poor performance due to this small-sample size. For this reason, in order to ascertain that there is no cointegrating relationship between the real exchange rate and this set of fundamentals in the Asian developing-country group, we test all possible subsets of the real exchange rate and its fundamentals by dropping one fundamental at a time until a cointegrated relation is found. We find that there exists a cointegrating relationship between the real exchange rate and its three fundamentals - $(y_{it}^* - y_{it}), (gov_{it}^* - gov_{it}), open_{it}$ - in the group of Asian developing countries. See Table 4 for the results of these tests.

Table 4: Cointegration tests for the real exchange rate relationship in the Asian developing group

Variables	Panel ADF	Group ADF
$q_{it}, (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it}$	0.01	0.05

See Table 2.

In summary, the final choice of the cointegrating relationship between real exchange rates and its fundamentals found in each country group can be summarized in Table 5. In the case of Asian developing countries, this finding corresponds with our results of the cointegrating tests of the developing-country group in the first chapter, and this is in line with controversy regarding real exchange rate movements in fast-growing Asian countries. That is, our results imply that there is no long-run (cointegrating) relationship between real exchange rates and productivity growth in traded sectors in these countries. This is consistent with the actual situation, as rapidly growing countries in Asia such as China and India do not seem to experience real exchange rate appreciation.

Table 5: Cointegrating relationships between the real exchange rate and its fundamentals

Country groups	The collection of cointegrated variables
Developed countries	
European countries	$q_{it}, (x^*_{it} - x_{it}), (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it}$
Other developed countries	$q_{it}, (x^*_{it} - x_{it}), (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it}$
Developing countries	
Asian developing countries	$q_{it}, (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it}$
Non-Asian developing countries	$q_{it}, (x^*_{it} - x_{it}), (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it}$

4.4 Estimating and interpreting the cointegrating vector

Section 4.3 found evidence for the presence of long-run cointegrating relationships in each group of countries, and in this section we estimate the cointegrating vectors and examine their signs. Following the literature, there are three common approaches used to estimate cointegrating vectors in panel data: ordinary least squares (OLS), fully modified OLS (FMOLS) suggested by Pedroni (1996, 2000), and dynamic OLS (DOLS) proposed by Kao and Chiang (1999) and Mark and Sul (2003). In general, these approaches allow for heterogeneous short-run dynamics, and pool the information along the within-dimension or along the between-dimension, depending on the long-run hypothesis of interest.

It is well known that the OLS estimator is typically biased in panel data settings, but although the FMOLS estimator was created to solve the bias of OLS estimators by using a complicated estimation technique to account for the endogeneity and serial correlation, Kao and Chiang (1999) show that it is still biased in finite samples and does not improve over the OLS estimator in general cases. Kao and Chiang (1999) and Pedroni

(2001) also find that the DOLS estimator has smaller size distortions and outperforms the others in finite samples. We decide to use the panel DOLS estimator in our study, for these reasons.

There are two techniques for estimating DOLS (1) between-dimension estimation and (2) within-dimension estimation. In this chapter, we choose to apply the within-dimension estimators as suggested by Mark and Sul (2003). This differs from the estimation method used in the first chapter, because although the between-dimension estimators have comparative advantages in terms of greater flexibility in the presence of heterogeneous cointegrating vectors, they are still applying time series estimation to short spans of data. On the other hand, the within-dimension estimators are somewhat restrictive since cointegrating vectors are supposed to be homogeneous across cross-sectional units. Nevertheless, given that we divide our set of countries into four groups, each of which includes country members with similar features, allowing for heterogeneity across countries through heterogeneous short-run dynamics, country-specific fixed effects, and allowing for a limited degree of cross-sectional dependence through time-specific effects seem sufficient to capture the heterogeneity in the country members of each country group[18]. With the within-dimension technique, we can also obtain more precise point estimates of the cointegrating vectors due to the improvement of finite-sample estimation by using the panel .

We focus on estimating a cointegrating vector of the relationship (5) as the main purpose of our study is to examine real exchange rate behavior in response to shocks in its fundamentals. The DOLS regression for each country group is given by

$$q_{it} = d_i + \theta_t + \beta' fd_{it} + \delta'_i sd_{it} + u_{it} \tag{34}$$

where $i = 1, 2, ..., N_j$ and N_j is the number of countries in group j, $fd_{it} = ((x^*_{it} - x_{it}), (y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it})'$ in the case of all groups with the exception of the Asian developing-country group for which $fd_{it} = ((y^*_{it} - y_{it}), (gov^*_{it} - gov_{it}), open_{it})'$; $sd_{it} = (\Delta fd'_{i,t-1}, \Delta fd'_{i,t}, \Delta fd'_{i,t+1})'$; d_i is a country-specific effect; θ_t is a common time-specific factor which is used to capture some forms of cross-sectional dependence across country members, as the asymptotic distribution theory requires the cross-sectional independence. We focus on the estimate of $(1, -\beta')$ which is a cointegrating vector between q_{it} and fd_{it}: we allow this estimate to be identical across individual countries in each group. The coefficients β' capture the long-run impact on q_{it} to long-run changes in the fd_{it} variables. It is obvious that DOLS uses the past and future values of Δfd_{it} to control for the endogeneity.

[18] In the first chapter, we separated countries into two groups: developed-country group and developing-country group. Each group consisted of a large set of country members with different characteristics. As a result, it is likely that the within-dimension technique cannot capture enough heterogeneity, and the between-dimension technique is more reasonable eventhough we lose efficiency in estimation.

In order to estimate a cointegrating vector from equation (34), we start with regressing q_{it} and fd_{it} on $(1, s\acute{d}_{it})^{\acute{}}$ for each country and collect the residuals, called q^{+}_{it} and fd^{+}_{it}, from these regressions. Time-specific dummy variables are also included in the regression of Brazil, Chile, and Argentina to control for irregular behaviors of macroeconomic series in the period of Latin American debt crisis; specifically, the time-specific dummy variables for Brazil, Chile, and Argentina cover the period 1980-1988, 1980-1988, and 1980-1990, respectively. To illustrate, see Figure C1 in Appendix C. Then we can obtain a cointegrating vector by estimating the following regression:

$$q^{+}_{it} = \acute{\beta} fd^{+}_{it} + \theta_t + u_{it}. \tag{35}$$

We include common time dummy variables into the regression in order to account for cross-sectional dependence across countries in panels and obtain residuals with no cross-sectional correlation. Table 6 reports the estimated cointegrating vector for the relationship (5) in each country group[19].

Table 6: The panel DOLS estimates of a cointegrating vector for each country group

Variables	Europe	Other developed	Asian Developing	Non-Asian developing
$x^*_{it} - x_{it}$	-0.439^{**}	-0.247^{**}	-	0.119
	(0.086)	(0.065)	-	(0.120)
$y^*_{it} - y_{it}$	0.145^{**}	-0.110^{**}	-0.539^{**}	0.302
	(0.049)	(0.043)	(0.074)	(0.212)
$open_{it}$	-0.137	0.685^{**}	0.071	0.048
	(0.120)	(0.065)	(0.056)	(0.159)
$gov^*_{it} - gov_{it}$	0.683^{**}	-0.213	0.276^{**}	0.425^{**}
	(0.196)	(0.153)	(0.097)	(0.136)
R^2	0.520	0.539	0.705	0.345
F-statistics	3.839^{**}	5.267^{**}	15.064^{**}	1.540^{*}

Note: 1) *, ** indicate 5%, 1% significance levels, respectively. Standard errors are reported in parentheses.
2) Due to the short span of data, we use one lag and one lead of independent variables in differences to account for endogeneity.

We see that the set of real exchange rate fundamentals in our study is larger than that suggested by the BSH, but smaller than that used in the Edwards model. In particular, not only do real exchange rates have a long-run comovement with the relative productivity gain in the traded sector, but also with the relative growth rate of real GDP, the degree of openness in the economy, and government consumption.

According to Table 6, most of the statistically significant results of the relationship between real exchange rates and its fundamentals in our country groups correspond to our expectation. However, two issues

[19] Note that the U.S. data set is not included when estimating a cointegrating vector for developed-country group. This is because in estimating region-specific VECM for this group in the next section, we exclude U.S. as it is likely that the weakly exogeneity assumption of U.S. foreign variables is violated. For the Fisher Equation and UIP condition, the estimates of cointegrating vectors are shown in Tables B5.1 and B5.3, Appendix B.

are worth mentioning. Firstly, the estimated cointegrating vectors show that for developed countries, a productivity improvement in traded sectors relative to nontraded sectors is associated with a real depreciation, after controlling for the growth rate of real GDP, the degree of openness in the economy, and the government consumption share. Secondly, there is a significant difference between the results of European and other-developed country groups with respect to the relationship between the real exchange rate and economic growth. Specifically, rapid economic growth in the European country group is associated with a real appreciation, whereas it is associated with a real depreciation in the other-developed country group, after controlling for the productivity gain in traded sector, the degree of openness in the economy, and the government consumption share.

We note that although the estimates of cointegrating vectors can represent the relationship between the real exchange rate and its fundamentals, these estimates cannot imply what the direction of causation among them is. Therefore our next task is to examine what the direction of causation are, how fast shocks to these fundamentals have impacts on long-run real exchange rates and other macroeconomic variables, how large these impacts are, and how long those variables, especially real exchange rates, take to return to their long-run equilibrium level.

4.5 Impulse Response Analysis

The vector error correction model (VECM) in each country group comprises seven endogenous/domestic variables and six exogenous/country-specific foreign variables. They enter the model as the real effective exchange rate (q_{it}), the productivity in the traded sector relative to the nontraded sector (x_{it}, x^*_{it}), real GDP (y_{it}, y^*_{it}), the government consumption share (gov_{it}, gov^*_{it}), the degree of openness in the economy ($open_{it}$), the nominal short-term interest rate (si_{it}, si^*_{it}), the inflation rate (π_{it}, π^*_{it}) and oil prices (oil_t). This set of variables is smaller than the set of variables at the beginning because we have dropped some variables which have no long-run cointegrating relationship with the real exchange rate for reasons of parsimony. Therefore for each country group examined, we estimate a VECM with the following endogenous and exogenous variables,

$$\Delta g_{i,t} = d_i - \alpha \acute{\hat{\beta}} z_{i,t-1} - \Phi_2 \Delta g_{i,t-1} + \Psi_0 \Delta g^*_{i,t} + u_{i,t},$$

where

$$\begin{aligned} g_{it} &= [q_{it}, x_{it}, y_{it}, gov_{it}, open_{it}, si_{it}, \pi_{it}]', \\ g^*_{it} &= [x^*_{it}, y^*_{it}, gov^*_{it}, si^*_{it}, \pi^*_{it}, oil_t]', \text{ and} \\ z_{it} &= \left(\acute{g}_{it}, g^{*}_{it}\right)' \end{aligned}$$

for $i = 1, 2, ..., N_j$ and N_j is the number of countries in group j. For estimating the VECM in each country group, we impose cointegrating restrictions $\widehat{\beta}_j$ computed by using panel DOLS estimation or suggested by economic theory. We also include country-specific dummy variables d_i, to account for country-specific fixed effects. Note that when we test long-run relationship (5), most variables are in differentials (differences between domestic and foreign variables); however, we separate them out into individual domestic variables and foreign variables when estimating the VECM.

As the weak exogeneity of foreign variables and oil prices is an important condition for obtaining consistent parameters from estimation, we exclude the USA from the model of the other-developed country group to avoid violation of this condition. Note that the results for the weak exogeneity tests on most of the country-specific foreign variables and oil prices fail to reject the null hypothesis of weak exogeneity at the 5% level, satisfying the presence of weak exogeneity of country-specific foreign variables (See Table B7, Appendix B). One exception is that weak exogeneity of foreign inflation rates in the non Asian developing-country group was rejected at the 5% significance level. However, this test result seems not to be convincing and plausible so we still treat foreign inflation rates as weakly exogenous in the non Asian developing-country model. Moreover, as mentioned before, we include time-specific dummy variables for Brazil, Chile, and Argentina to account for debt crisis in the early 1980s when estimating the model.

Figures C2-C17 in Appendix C show impulse responses to four structural innovations - productivity growth, trade liberalization, contractionary monetary policy, and government spending shocks - for four country groups, i.e. European group, other developed-country group, Asian developing-country group, and non Asian developing-country group, based on sign restrictions with a penalty-function approach. Shocks are normalized to the magnitude of one standard deviation in size. Each plot comprises of the solid line, which represents the median of the posterior distribution of impulse responses, and the dashed lines, which represent as the 16 and 84 percent quantiles of the posterior distribution, corresponding to the one standard deviation error bands, assuming that variables follow a normality distribution. We show the impulse responses of seven endogenous variables over nine years after the shocks.

4.5.1 Real Exchange Rate Responses

In our study, we focus on the movements of real exchange rates in response to each underlying shock in real exchange rate determinants, a monetary policy shock, and a fiscal policy shock. We summarize the responses of real exchange rates to the shocks in Table B8, Appendix B. Recall that the exchange rate is defined so that it declines as the value of home country currency increases.

Productivity Improvement Shock Figure 1 shows the responses of real exchange rates to a productivity improvement shock in each country group. As shown in Figures C2-C5, Appendix C, one can clearly see that although no sign restriction is placed on the responses of productivity differential and real output after one year, their responses in all country groups are persistent.

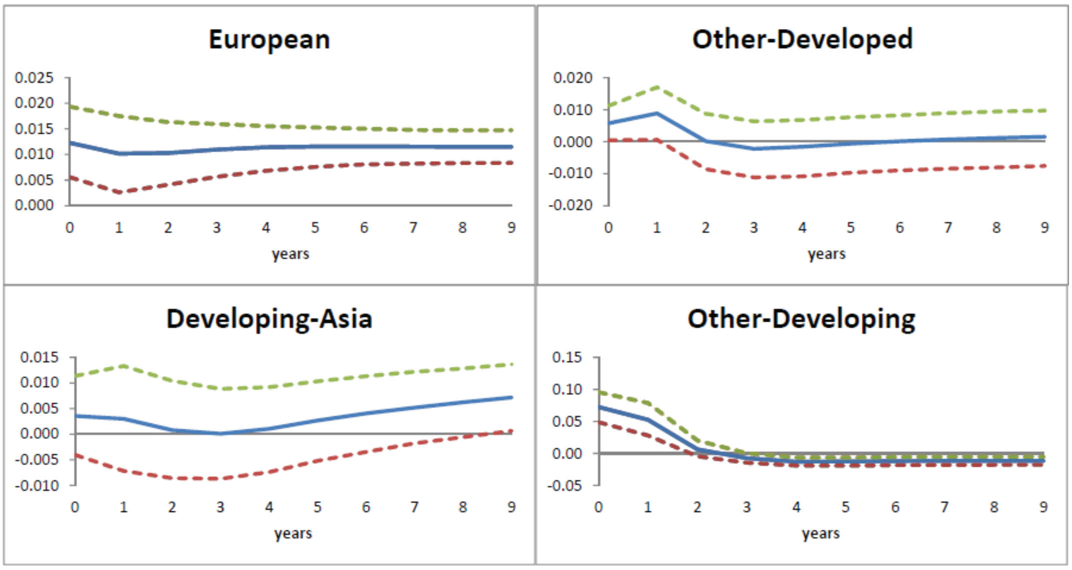

Figure 1: Impulse responses of real exchange rates to a productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function.

According to the traditional view, if productivity in the traded sector in one country grows faster than that in other countries (given that productivity growth in the nontraded sector for all countries is slow), its real exchange rate tends to appreciate. This is consistent with our finding for the non Asian developing-country group, i.e. after a shock on productivity improvement in the traded sector, its real exchange rate tends to persistently appreciate after two years. However, our results show that the productivity shock has no effects on real exchange rates in the Asian developing group. This corresponds to the actual situation in Asia as fast economic growth in some countries such as China and India does not seem to be followed by a real appreciation.

Moreover, we find that the shock leads to a long-run real depreciation in the European country group. These surprising results may be explained by using the Edwards model. That is, although productivity growth induces an income effect and then causes a real appreciation, it is possible that productivity improvement can lead to a real depreciation if the supply effects dominate the demand effects. In particular, productivity growth in the traded sector implies a rise in the marginal product of labour; thus increasing the real wage. In order to maintain the same level of workers in the nontraded sector, wages in this sector need to be higher. It is likely that a small increase in wage of the nontraded sector attracts some workers from the traded sector into the nontraded sector. Alternatively, an increase in productivity in traded sectors relative to nontraded sectors can be a result of relatively more factor augmenting progress in nontraded sectors than in traded sectors, by construction[20]. As a result, more workers or an increase in a factor of production in nontraded sectors can lead to excess supply of nontraded goods, causing a fall in their prices. This leads to a real depreciation.

Moreover, these results may be explained by a simply small open economy model, developed by Devarajan, Lewis, and Robinson (1990). According to the Devarajan model, productivity progress in the traded sector shifts up the productivity possibility frontier (PPF), increasing real output, and thereby expanding the consumption possibility frontier (CPF). The income effect will generate higher demands in all goods, inducing higher prices. However, the effects on real exchange rates also depend on preferences between traded and nontraded goods, or in other word, the elasticity of substitution between imports and nontraded goods. If a rise in income is accompanied by higher demand of traded goods i.e. imports, relative to the demand of nontraded goods, then this can lead to a fall in relative demand of nontraded goods and their prices, and thereby depreciates real exchange rates.

Trade-Liberalization Shock The impulse responses of real exchange rates to a trade liberalization shock are displayed in Figure 2. According to the traditional view and the Edwards model, trade liberalization or a decline in trade restrictions, i.e. import tariffs, increases foreign goods in the domestic country and hence generates a trade account deficit. In order to restore equilibrium, real exchange rates tend to depreciate in the long run. Interestingly, our findings in all country groups are in line with this view. That is, trade liberalization shocks in all country groups are followed by a real depreciation.

[20]As mentioned before, the productivity variable is constructed as the value added divided by the number of workers in that sector. Thus when the number of workers in the nontraded sector increases given the same level of the value added, the productivity in the nontraded sector will decrease, by construction. This leads to an increase in the productivity differential between traded and nontraded sectors.

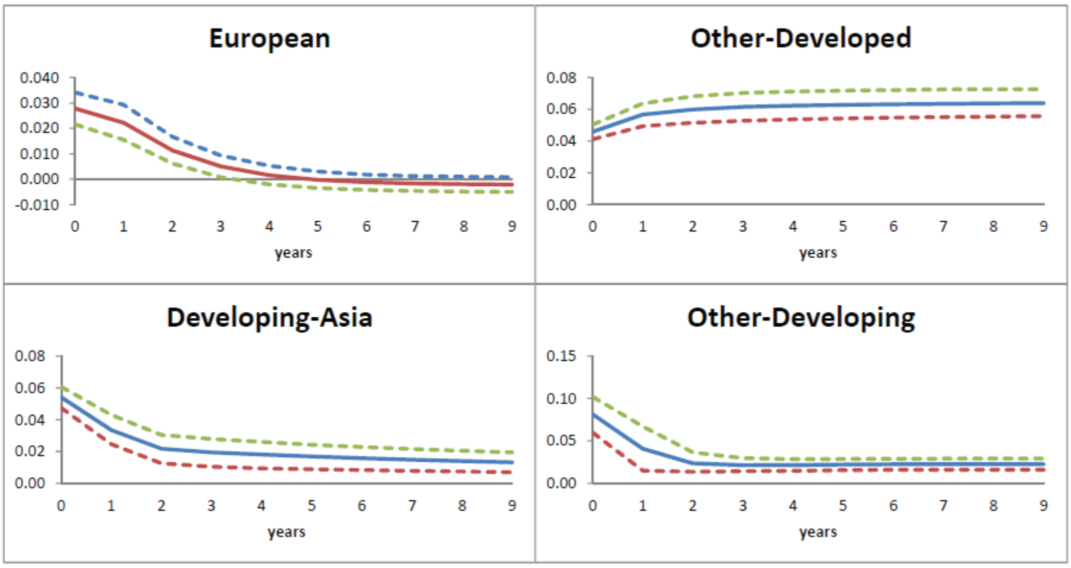

Figure 2: Impulse responses of real exchange rates to a trade-liberalization shock of one standard deviation in size, using sign restrictions with the penalty function.

Contractionary Monetary Policy Shock A contractionary monetary policy shrinks the total money supply in the economy or raises interest rates in order to reduce inflation and maintain stability. The impulse responses of real exchange rates to the contractionary monetary policy shock are displayed in Figure 3. Note that the price puzzle does not occur in our study, by construction. The results show that the contractionary monetary policy shock has no permanent effects on the real exchange rate in any of the country groups, consistent with the long-run neutrality of monetary policy.

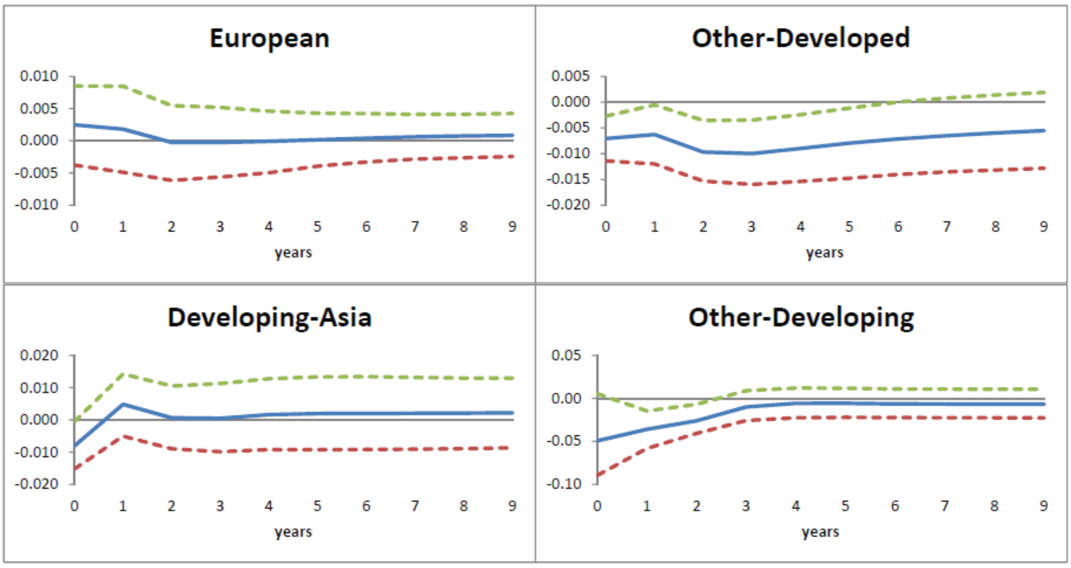

Figure 3: Impulse responses of real exchange rates to a contractionary monetary policy shock of one standard deviation in size, using sign restrictions with the penalty function.

Following the literature, an increase in interest rates causes net inflows on the capital account, boosting the supply of foreign currencies. As a result, the price of foreign currency falls and this can lead to an appreciation of domestic currency in the short run. Our finding shows that after the shock, real exchange rates in the other-developed, and the other-developing country groups move in the same direction as we expect; i.e. higher interest rates are associated by an appreciation in the short run. However, we find that the monetary policy shock does not have any short-run effect on real exchange rates in the European and the developing-Asia groups.

Government Spending Shock According to existing literature, government consumption has a significant influence on long-run real exchange rate movements. The impulse responses of real exchange rates to this shock are presented in Figure 4.

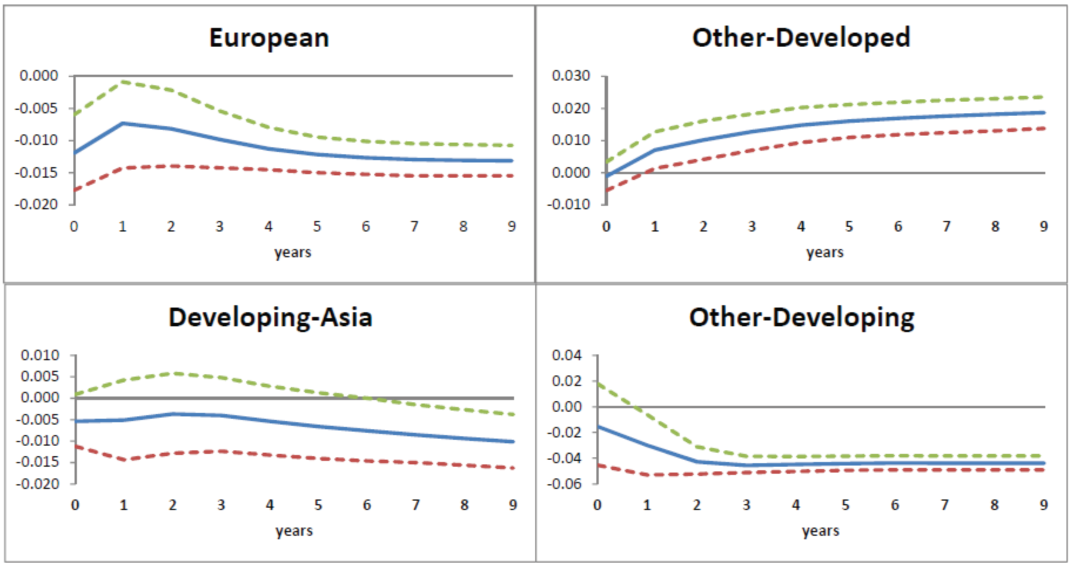

Figure 4: Impulse responses of real exchange rates to a government spending shock of one standard deviation in size, using sign restrictions with the penalty function.

Following the traditional view, an increase in government consumption is regarded as a rise in the relative demand for nontraded goods, thereby increasing their prices and leading to a real appreciation. For the European, the developing-Asia, and the non Asian developing-country groups, the responses show that the government spending shock leads to an appreciation in real exchange rates in the long run, which is in line with the prediction of the theoretical exchange rate model and the existing literature such as Chinn (1999), Galstyan and Lane (2009). For the developing-Asia group, the responses of real exchange rates to an increase in government consumption share are very slow, i.e. the shock starts to effect the movement of real exchange rates after 6 years.

However, we find that the shock in the other-developed country group is followed by a real depreciation in the long run. This surprising result may be explained by the Edwards model. That is, in this country group, if a tax increase is associated with a shock in government spending and if the corresponding decline in household income has more dominant effects than the initial increase in government spending, this can tend to a decline in demand and consumption in all goods, causing a fall in the price of nontraded goods and thus a real depreciation.

Overall, most of our results are in line with international macroeconomic theories and our expectation. In particular, for the two groups of developing countries examined, all responses of real exchange rates to the four shocks in real exchange rate fundamentals and a change in policy are consistent with traditional views and our expectation. Only two exceptions are found: the responses of real exchange rates to the productivity shock in the European country group, and the responses to government spending shock in the other developed-country group.

4.5.2 Other interesting responses

In addition to the responses of real exchange rates to the shocks, there are some interesting results that we obtain from using sign restrictions and the penalty-function approach as follows.

Productivity Improvement Shock A shock to productivity growth in the traded sector has a negative impact on the government consumption share in almost all country groups. This finding can imply that a positive income effect after the shock leads to an increase in individual consumption more than a rise in government consumption. It seems reasonable because the consumption of individuals is flexible and mainly depends on real income, while government consumption is a policy variable and so it is rigid and hard to change. The responses are shown in Figures C2-C5, Appendix C.

Trade Liberalization Shock Interestingly, a trade liberalization shock is followed by inflation in all country groups. This may be a consequence of an income/welfare effect. In particular, a reduction of an import tariff stimulates a rise in demands for all goods and thus their prices. Moreover, we also find that for the two groups of developed countries, trade liberalization can help enhance productivity in the traded sectors. This is in line with traditional trade theory as it is known that trade liberalization is one factor that can help improve productivity performance through an efficient allocation of resources between countries. Also, a trade reform stimulates the incentive for domestic firms to compete or seek new markets overseas, encouraging firms to adopt new ideas and invest in new technology. These can be translated into productivity gains. The responses are shown in Figures C6-C9, Appendix C.

Contractionary Monetary Policy Shock In addition to no long-run response of real exchange rates to a surprise change in nominal interest rates, a monetary policy shock has no long-run effect on real GDP in most country groups, as expected. These effects are consistent to the idea that money is neutral. The results are displayed in Figures C10-C13, Appendix C.

Given long-run neutrality of monetary, the demand and supply for real money balances are unchanged after the shock. However, a rise in interest rates tends to raise domestic saving and reduce domestic investment (as the opportunity cost of the funds is higher). As a result, lower nominal money supply in the economy causes a rise in the price of money or, equivalently, lower prices of goods, i.e. a decline in inflation. This corresponds to our results for all country groups.

However, the neutrality of money does not completely hold as the monetary policy shock has some impacts on other real variables. For the other developed-country and the Asian developing-country groups, productivity in traded sectors falls after the shock. A plausible explanation is that a decline in domestic investment after the shock might be followed by a fall of the steady-state capital to labour ratio, and this might cause a drop in productivity in traded sectors.

Furthermore, the results reveal that a monetary policy shock has an negative impact on real GDP in the non Asian developing-country group. This might be explained by nominal wage rigidity. That is, a reduction of prices after the shock causes an increase in the real wage, so less workers are hired, thereby leading to a fall in real output. Under the market clearing condition, a rise in real wages also implies a productivity improvement. Our result confirms this idea as productivity in this country group increases after the shock.

Government Spending Shock The responses for a government spending shock are shown in Figures C14-C17, Appendix C. As seen in the literature, the government spending shock might be considered as a fiscal policy shock. Specifically, a higher government consumption share can indicate expansionary fiscal policy. In the European group, an increase in government consumption boosts the aggregate demand for goods and thus causes inflation. However, the implications of this do not seem to correspond with Keynesian economics, but rather to the ideas of Classical and Neoclassical economists who debate the effectiveness of fiscal policy for stimulating the economy. That is, for running a budget deficit, government must borrow funds from the public or from overseas. This causes a higher demand for money and thereby a rise in interest rates. Higher interest rates attract a foreign capital inflow, inducing a surplus in the capital account and thus causing a real appreciation. Once the domestic currency appreciates, net exports fall and finally this can lead to a decline in real GDP in the economy.

However, fiscal policy actually relies on two main components - government spending and government revenue. Therefore, using only government spending shocks as representative of fiscal policy can lead to misleading conclusions. This can be seen from counter-intuitive responses in the Asian-developing group i.e. an expansionary fiscal policy shock causes a fall in interest rates and inflation. This may be because an increase in government spending share in this country group is a result of the contractionary fiscal policy shock, not expansionary fiscal shock, or in other words, a rise in the government consumption share may be a consequence of higher government revenue i.e. an increase in tax revenue. As mentioned before, real income is a main determinant of private consumption and individual consumption is very responsive to a change in real income. Consequently, a fall in private consumption due to higher tax may lead to a rise in the government consumption share. Also, a negative income effect causes deflation in this country group as shown in our results.

4.6 Forecast Error Variance Decomposition Analysis

We turn to a forecast error variance decomposition with a sign restriction approach. This analysis answers the question of how much of the variances in real exchange rates over the sample period can be explained by the four shocks. Table B9 in Appendix B shows, for each country group examined, the percentage of the variances of real exchange rates accounted for by the productivity improvement shock, the trade liberalization shock, the contractionary monetary policy shock, and the expansionary government spending shock.

We find that these four shocks can explain the large proportion of the forecast error variance of real exchange rates in all country groups examined. In detail, for nine years after the shocks, the peaks of the contribution of total shocks are around 40%, 71%, 35%, and 36% in the European, the developed-country, the Asian developing-country, and the non Asian developing-country groups, respectively. We also find that the fractions of variation in real exchange rates explained by the shocks are larger at longer horizon. This may be because the responses of real exchange rates to the shocks are very slow.

Interestingly, the results suggest that for almost all country groups examined, the trade liberalization shock is an important factor in explaining the forecast error variance of real exchange rates, especially in the other developed-country group i.e. for this country group, the trade liberalization shock explains more than 55% on average of the real exchange rate forecast error variance. Also, for the European and the Asian developing-country groups, such a shock contributes for almost 15% and 27% of real exchange rate forecast error variance, respectively. The exception is the non Asian developing-country group; the trade liberalization shock accounts for only 7% of real exchange rate forecast error variance. In particular, among the four shocks examined, we find that the trade liberalization shock contributes the most to real exchange rate error variance for almost all country groups except the non Asian developing-country group for which the government spending shock seems to be the most important factor among these four shocks. However, the trade liberalization shock is still the second most important factor in explaining the forecast error variance of the real exchange rate in the non Asain developing-country group. This implies that for the European, the developed-country, the Asian developing-country groups, international trade policy can be a more efficient instrument in determining real exchange rate behavior compared to other policies, whereas fiscal policy is more effective than the others in the non Asian developing-country group.

5 Conclusion

The existing literature on the relationships between the real exchange rate and its determinants for developed countries is extensive. However, the evidence on these relationships for developing countries has been largely unexplored. The main purpose of this paper is to examine the role of a parsimonious set of exchange rate fundamentals on real exchange rate behavior in thirty countries that include sixteen developed countries and fourteen developing countries over the period 1970-2008. The selection of exchange rate fundamentals in our study is based on the BSH and Edwards models. We divide our country set into four groups (European, developed-country, Asian developing-country, and non Asian developing-country groups) according to the region and the level of economic development, and then apply panel estimation techniques.

For our study, we construct a sectoral productivity variable as a proxy for the Balassa-Samuelson effect, using a novel approach for the classification of industries into traded and nontraded sectors, introduced in the first chapter. The long-run cointegrating relationship between the real exchange rate and its determinants is examined by using the Pedroni (2000) panel cointegration tests together with the parsimonious approach of Davidson (1998). Panel dynamic ordinary least squares (DOLS) estimation suggested by Mark and Sul (2003) is then employed to estimate the panel cointegrating vectors. We employ a new method based on sign restrictions and a penalty-function approach to identify the effects of four shocks - productivity improvement, trade liberalization, contractionary monetary policy, and government spending shocks - on real exchange rate behavior. This approach requires only a set of a priori theoretical restrictions that are often used implicitly by researchers, and leaves the question of interest open. It is also superior to other identification methods as it generates results that are robust to the choice of reordering variables or selecting a Cholesky decomposition. The most innovative aspect of our study is that we adjust the standard sign restriction approach based on a structural vector autoregression model (SVAR) to a sign restriction approach based on a Bayesian structural vector error correction model that models panels of data. For all the country groups examined, impulse responses and variance decomposition are considered.

Our main finding is that the traded-nontraded productivity differential, real GDP growth, government consumption share, and the degree of openness in the economy have a long-run cointegrating relationship with the real exchange rate for almost all country groups, with the exception of the Asian developing-country group in which the sectoral productivity differential seems to have no cointegrating relationship with the real exchange rate. Most of our impulse response analysis confirms the results expected by economic theories and found in the empirical literature so far. Firstly the results in all country groups examined suggest that trade liberalization can lead to a long-run real depreciation. Secondly, we find strong evidence that a rise in government spending generates a real appreciation in the long run in almost all country groups. Thirdly, our finding shows that a contractionary monetary policy shock has no long-run impacts on real exchange rate movements in any of the country groups examined, although for some country groups, the shock may influence the movements of the real exchange rate in the short run. This corresponds to the long-run neutrality of monetary policy.

Moreover, the shock on productivity improvement leads to a real appreciation in the non Asian developing-country group as expected. However, two exceptions are, of course, significant differences. Firstly, we find that the productivity shock has no effects on real exchange rates of the Asia-developing country group. This corresponds to the controversy regarding real exchange rate behavior among fast-growing countries in Asia. Secondly, for the European group, productivity growth in traded sectors relative to nontraded sectors leads to a long-run real depreciation. One plausible explanation is that higher wages in traded sectors due to productivity growth are followed by a rise of wages in nontraded sectors. This can lead to factor-augmenting progress in nontraded sectors, thereby creating a positive supply effect in nontraded goods. If this induces excess supply in nontraded goods, implying a fall in their prices, then the real exchange rate will depreciate.

Variance decomposition suggests that the productivity improvement, the trade liberalization, monetary policy, and government spending are important for explaining the dynamic movement of the real exchange rate. Our results reveal that trade liberalization seems to be the most important factor for the movement of real exchange rates in almost all country groups, as opposed to monetary policy or fiscal policy. However, the exception is found in the non Asian developing-country group for which fiscal policy via a change in government spending has a more influence on the real exchange rate than the others.

Due to global integration, it is possible that external shocks around the world can also influence the movement of real exchange rates. Therefore, an examination of this issue within a global context might further help in explaining real exchange rate behavior. We have left this point for the third chapter.

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A Appendix

A.1 Countries

The data set consists of an unbalanced panel data from 1970 to 2008 for thirty countries; five countries that joined the European monetary union in 1999 are grouped together into a single Euro area economy. We use the same country set as in the first chapter (excluding Myanmar, Brunei Darussalam, and Vietnam) in order to capture all possible influence of external shocks on the movements of real exchange rates, and to obtain appropriate country-specific foreign variables. We classify our country set into four groups according to the level of economic development and regions - European countries, other developed countries, developing countries in Asia, and other non-Asian developing countries - in order to reduce the limitations associated with small sample size and take advantage of panel techniques. Table A1 provides the details of country members in each country group.

Table A1: List of countries in each group

Country Grouping	Members
Developed countries	
European countries (9 countries)	Euro Area (Germany, France, Italy, Spain, Netherlands), U.K., Norway, Sweden, Switzerland
Other developed countries (7 countries)	Australia, New Zealand, Canada, USA Korea, Japan, Singapore
Developing countries	
Developing Asian countries (8 countries)	China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan, Sri Lanka
Other developing countries (6 countries)	Brazil, Mexico, Chile, Argentina, South Africa, Turkey

Note: we classify countries into developed and developing countries according to the International Monetary Fund's World Economic Outlook Report, April 2010.

A.2 Data Sources

A.2.1 Consumer Price Indices (CPI)

The data source for all countries was the World Bank, World Development Indicators (WDI). For Germany, OECD Main Economic Indicators database (MEI) data were used for 1961-1990. Data from the IMF's

International Financial Statistics (IFS) were used to complete gaps in the WDI series for China.

A.2.2 Nominal Exchange Rate

The data source for all countries was the WDI. With the exception in Euro area, the exchange rate series were monthly exchange rate series constructed by Anderson et al (2011). We transformed this into annual series by using simple average.

A.2.3 GDP/Value Added

The data source for all countries was the National Accounts Main Aggregates database, compiled by The United Nations (UN). We used GDP measured in current and constant 1990 local currency units, classified by economic activity (International Standard Industrial Classification: ISIC 3) into seven categories.

A.2.4 Short-Term Interest Rate

The data source for almost all countries, except Australia and New Zealand, was the IFS. The data source for Australia and New Zealand was MEI. For most countries, the IFS money market rate series were used as the short-term interest rate with the exception as follows. The IFS deposit rate series were used for China, Chile, and Netherland. The IFS treasury bill rate series were used for France, UK, Canada, Mexico, and Sweden. Some interest rate series, i.e. the OECD short-term interest rate, IFS treasury bill rate, and IFS deposit rate for some countries where IFS money market rate series were available in recent period, were spliced at the end in order to be compatible with IFS money market rate for other countries. The detail for interest rate series in each country is available upon request.

A.2.5 Export Value Index and Import Value Index

The data source for Australia, New Zealand, Germany, France, United Kingdom, Italy, Spain, Canada, Brazil, Norway, United States, Japan, Korea, Thailand, India, and Pakistan was the IFS. For the rest, the series were taken from the WDI.

A.2.6 Government Consumption (% of GDP)

The government consumption share series for all countries were from the WDI. Data from the IFS were used to complete gaps in the WDI series for New Zealand and Germany.

A.2.7 Exports and Imports (% of GDP)

The data source for all countries was the WDI. Data from the IFS were used to complete gaps in the WDI series for Singapore.

A.2.8 Oil Price Index

Oil prices we used are averages of Brent Crude series from Datastream. Data from Energy Information Administration of USA were used to fill the gaps.

B Table Appendix

Table B1: Trade fixed weights based on direction of trade statistics

	China	Japan	Korea	Malaysia	Indonesia	Philippines	Thailand	Singapore	India	Pakistan	Sri Lanka	Australia	N. Zealand	U.K.	EU.	Canada	Brazil	Mexico	Chile	Norway	Sweden	Turkey	Switzerland	Argentina	S. Africa	U.S.
China	0.00	0.22	0.24	0.11	0.10	0.09	0.11	0.12	0.12	0.14	0.06	0.14	0.10	0.05	0.11	0.05	0.09	0.05	0.12	0.03	0.04	0.07	0.03	0.10	0.09	0.14
Japan	0.20	0.00	0.18	0.15	0.21	0.20	0.23	0.10	0.05	0.06	0.05	0.17	0.12	0.04	0.06	0.03	0.05	0.03	0.10	0.02	0.03	0.03	0.04	0.02	0.11	0.10
Korea	0.11	0.08	0.00	0.05	0.07	0.06	0.04	0.05	0.04	0.04	0.03	0.06	0.04	0.01	0.03	0.01	0.03	0.02	0.05	0.01	0.01	0.03	0.01	0.01	0.03	0.04
Malaysia	0.03	0.03	0.03	0.00	0.05	0.06	0.07	0.18	0.03	0.04	0.03	0.03	0.03	0.01	0.01	0.00	0.01	0.01	0.00	0.00	0.00	0.01	0.00	0.01	0.01	0.02
Indonesia	0.02	0.04	0.03	0.04	0.00	0.02	0.04	0.09	0.03	0.03	0.02	0.03	0.02	0.00	0.01	0.00	0.01	0.00	0.01	0.00	0.00	0.01	0.00	0.01	0.01	0.01
Philippines	0.02	0.02	0.02	0.02	0.01	0.00	0.02	0.03	0.01	0.00	0.00	0.01	0.01	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.01
Thailand	0.02	0.04	0.02	0.06	0.04	0.04	0.00	0.05	0.02	0.02	0.02	0.04	0.02	0.01	0.01	0.00	0.01	0.00	0.01	0.00	0.00	0.01	0.01	0.01	0.02	0.01
Singapore	0.03	0.03	0.03	0.16	0.13	0.10	0.07	0.00	0.06	0.03	0.07	0.05	0.03	0.02	0.02	0.00	0.01	0.01	0.00	0.01	0.01	0.00	0.01	0.00	0.01	0.02
India	0.02	0.01	0.02	0.03	0.04	0.01	0.02	0.03	0.00	0.04	0.19	0.03	0.01	0.02	0.02	0.00	0.01	0.00	0.02	0.00	0.01	0.01	0.01	0.02	0.03	0.01
Pakistan	0.00	0.00	0.00	0.00	0.01	0.00	0.00	0.00	0.01	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Sri Lanka	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.02	0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Australia	0.03	0.04	0.04	0.03	0.04	0.02	0.04	0.03	0.04	0.02	0.02	0.00	0.25	0.01	0.01	0.00	0.01	0.00	0.01	0.00	0.01	0.01	0.01	0.01	0.03	0.01
N. Zealand	0.00	0.01	0.00	0.01	0.01	0.01	0.01	0.01	0.00	0.00	0.01	0.06	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
U.K.	0.03	0.03	0.02	0.02	0.02	0.01	0.03	0.03	0.06	0.07	0.10	0.05	0.05	0.00	0.21	0.03	0.03	0.01	0.02	0.23	0.11	0.10	0.06	0.02	0.10	0.04
EU.	0.15	0.11	0.10	0.09	0.09	0.11	0.09	0.09	0.17	0.19	0.12	0.11	0.11	0.49	0.00	0.05	0.23	0.06	0.19	0.42	0.46	0.52	0.66	0.17	0.31	0.14
Canada	0.02	0.02	0.02	0.01	0.01	0.01	0.01	0.01	0.02	0.02	0.02	0.02	0.02	0.02	0.02	0.00	0.02	0.03	0.02	0.04	0.01	0.01	0.01	0.01	0.02	0.24
Brazil	0.02	0.01	0.01	0.00	0.01	0.01	0.02	0.01	0.01	0.01	0.00	0.01	0.00	0.01	0.02	0.01	0.00	0.01	0.09	0.01	0.01	0.01	0.01	0.30	0.02	0.02
Mexico	0.01	0.01	0.01	0.01	0.00	0.00	0.01	0.01	0.01	0.00	0.00	0.01	0.01	0.00	0.02	0.03	0.03	0.00	0.04	0.00	0.00	0.00	0.01	0.03	0.00	0.14
Chile	0.01	0.01	0.01	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.03	0.01	0.00	0.00	0.00	0.00	0.00	0.09	0.00	0.01
Norway	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.04	0.04	0.01	0.01	0.00	0.00	0.00	0.13	0.01	0.00	0.00	0.00	0.00
Sweden	0.01	0.01	0.00	0.00	0.01	0.00	0.01	0.00	0.01	0.01	0.01	0.01	0.01	0.03	0.05	0.00	0.01	0.00	0.01	0.12	0.00	0.02	0.01	0.01	0.01	0.01
Turkey	0.01	0.00	0.01	0.00	0.01	0.00	0.00	0.00	0.01	0.02	0.01	0.00	0.00	0.02	0.04	0.00	0.00	0.00	0.01	0.01	0.01	0.00	0.01	0.01	0.01	0.00
Switzerland	0.01	0.01	0.01	0.01	0.00	0.00	0.01	0.01	0.05	0.02	0.01	0.01	0.01	0.02	0.09	0.00	0.01	0.00	0.00	0.01	0.02	0.04	0.00	0.01	0.02	0.01
Argentina	0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.11	0.00	0.09	0.00	0.00	0.00	0.00	0.00	0.01	0.00
S Africa	0.01	0.01	0.01	0.00	0.01	0.00	0.01	0.00	0.03	0.01	0.00	0.01	0.01	0.02	0.02	0.00	0.01	0.00	0.00	0.00	0.01	0.01	0.01	0.01	0.00	0.01
U.S.	0.22	0.24	0.19	0.19	0.12	0.23	0.15	0.15	0.19	0.21	0.20	0.13	0.15	0.17	0.19	0.75	0.26	0.74	0.19	0.08	0.11	0.10	0.11	0.15	0.13	0.00

Note: Trade weights are computed as average shares of exports and imports over 2002-2008 displayed in columns by country such that a column, but not a row, sums to one.

Source: Direction of Trade Statistics, 2002-2008, IMF.

Table B2: Unit root test statistics for the level and first differences of domestic variables in each country

Countries/ Regions	q_{it}	Δq_{it}	x_{it}	Δx_{it}	rp_{it}	Δrp_{it}	r_{it}	Δr_{it}	tt_{it}	Δtt_{it}
Individual ADF-GLS unit root tests (P-Values)										
China	0.27	0.01**	0.93	0.02*	0.92	0.02*	0.01**	0.01**	0.14	0.02*
Malaysia	0.68	0.02*	0.24	0.01**	0.76	0.02*	0.01**	0.02*	0.01**	0.02*
Indonesia	0.57	0.02*	0.07	0.02*	0.29	0.02*	0.67	0.73	0.54	0.03*
Philippines	0.09	0.02*	0.12	0.01**	0.08	0.02*	0.01**	0.02*	0.30	0.01**
Thailand	0.45	0.02*	0.01**	0.02*	0.64	0.02*	0.07	0.02*	0.15	0.02*
Vietnam	0.59	0.31	0.53	0.02*	0.42	0.03*	0.88	0.16	0.72	0.01**
Brunei	0.27	0.01**	0.50	0.02*	0.40	0.02*	0.11	0.99	0.50	0.01**
India	0.72	0.02*	0.69	0.02*	0.33	0.02*	0.01**	0.02*	0.03*	0.02*
Pakistan	0.71	0.02*	0.25	0.03*	0.39	0.02*	0.01**	0.02*	0.19	0.02*
Sri Lanka	0.42	0.01**	0.09	0.01**	0.25	0.02*	0.01**	0.03*	0.42	0.02*
Brazil	0.62	0.02*	0.02*	0.03*	0.71	0.04*	0.06	0.02*	0.51	0.01**
Mexico	0.01**	0.02*	0.85	0.01**	0.76	0.02*	0.01**	0.02*	0.17	0.01**
Chile	0.27	0.02*	0.25	0.01**	0.06	0.23	0.01**	0.02*	0.14	0.01**
Turkey	0.35	0.51	0.51	0.02*	0.70	0.02*	0.01**	0.02*	0.74	0.02*
Argentina	0.17	0.02*	0.33	0.02*	0.94	0.01**	0.01**	0.02*	0.02*	0.01**
South Africa	0.56	0.02*	0.18	0.02*	0.79	0.02*	0.01**	0.02*	0.08	0.03*
Japan	0.55	0.02*	0.72	0.02*	0.39	0.02*	0.01**	0.03*	0.50	0.02*
Korea	0.01**	0.02*	0.68	0.38	0.43	0.02*	0.11	0.01**	0.70	0.01**
Singapore	0.49	0.02*	0.14	0.02*	0.64	0.02*	0.01**	0.20	0.23	0.01**
Australia	0.14	0.02*	0.84	0.02*	0.36	0.01**	0.10	0.02*	0.86	0.02*
New Zealand	0.04*	0.01**	0.58	0.02*	0.52	0.01**	0.22	0.71	0.12	0.02*
Euro	0.10	0.02*	0.52	0.02*	0.48	0.02*	0.11	0.03*	0.07	0.92
U.K.	0.32	0.02*	0.60	0.02*	0.01**	0.01**	0.03*	0.02*	0.09	0.02*
Canada	0.16	0.01**	0.98	0.02*	0.50	0.01**	0.05*	0.03*	0.41	0.02*
Norway	0.33	0.02*	0.02*	0.01**	0.67	0.02*	0.15	0.02*	0.78	0.01**
Sweden	0.93	0.02*	0.89	0.01**	0.93	0.01**	0.03*	0.03*	0.12	0.02*
Switzerland	0.53	0.02*	0.02*	0.02*	0.61	0.02*	0.02*	0.02*	0.17	0.01**
U.S.	0.21	0.02*	0.02*	0.03*	0.28	0.01**	0.04*	0.01**	0.71	0.02*
Panel CADF-GLS unit root tests (Test statistics)										
CIPS	-1.8	-5.2**	-2.4	-4.9**	-2.1	-5.0**	-2.7*	-4.9**	-2.2	-4.3**
Fisher	43.0	333.7**	68.2	442.0**	47.9	504.9**	128.5*	515.0**	43.2	364.8**
Inverse normal	-1.1	-15.8**	1.6	-16.7**	1.3	-18.3**	-4.9*	-17.1**	2.9	-14.2**

Note: 1) The tests for each variable are based on regressions that include an intercept and/or a linear trend.

2) For individual test, the numbers stand for p-values, while the numbers for panel test represent test statistics.

3) * and ** indicate significance at 5 and 1% levels, respectively. The critical values are computed using 20,000 simulations.

4) The lag length for each test is chosen using BIC with the maximum lag of 4.

Table B2 (continued)

Countries/ Regions	gov_{it}	Δgov_{it}	$open_{it}$	$\Delta open_{it}$	y_{it}	Δy_{it}	π_{it}	$\Delta \pi_{it}$	si_{it}	Δsi_{it}
Individual ADF-GLS unit root tests (P-Values)										
China	0.77	0.02*	0.49	0.01**	0.04*	0.01**	0.01**	0.01**	0.62	0.01**
Malaysia	0.39	0.02*	0.73	0.02*	0.24	0.02*	0.01**	0.02*	0.03*	0.02*
Indonesia	0.13	0.03*	0.41	0.01**	0.70	0.02*	0.02*	0.02*	0.01**	0.02*
Philippines	0.18	0.02*	0.62	0.02*	0.52	0.02*	0.01**	0.02*	0.10	0.02*
Thailand	0.04*	0.01**	0.49	0.02*	0.26	0.02*	0.01**	0.02*	0.23	0.02*
Vietnam	0.81	0.09	0.12	0.02*	0.34	0.03*	0.89	0.15	0.68	0.25
Brunei	0.91	0.01**	0.04*	0.03*	0.94	0.02*	0.02*	0.88	0.42	0.91
India	0.43	0.02*	0.92	0.04*	0.95	0.02*	0.02*	0.02*	0.01**	0.02*
Pakistan	0.06	0.03*	0.20	0.02*	0.36	0.02*	0.01**	0.02*	0.11	0.02*
Sri Lanka	0.40	0.03*	0.63	0.02*	0.74	0.02*	0.01**	0.03*	0.01**	0.02*
Brazil	0.01**	0.04*	0.78	0.02*	0.03*	0.02*	0.30	0.02*	0.22	0.02*
Mexico	0.51	0.02*	0.59	0.02*	0.50	0.03*	0.14	0.02*	0.45	0.01**
Chile	0.26	0.02*	0.36	0.02*	0.53	0.02*	0.01**	0.02*	0.49	0.50
Turkey	0.45	0.02*	0.01**	0.05*	0.96	0.02*	0.24	0.02*	0.45	0.02*
Argentina	0.55	0.01**	0.33	0.02*	0.66	0.01**	0.17	0.02*	0.02*	0.02*
South Africa	0.41	0.02*	0.19	0.02*	0.77	0.02*	0.25	0.02*	0.14	0.02*
Japan	0.25	0.01**	0.50	0.02*	0.61	0.10	0.13	0.02*	0.39	0.02*
Korea	0.41	0.02*	0.79	0.01**	0.98	0.01**	0.06	0.02*	0.61	0.02*
Singapore	0.03*	0.02*	0.27	0.01**	0.92	0.02*	0.02*	0.01**	0.08	0.03*
Australia	0.69	0.89	0.13	0.02*	0.20	0.77	0.20	0.02*	0.16	0.02*
New Zealand	0.92	0.24	0.30	0.02*	0.80	0.65	0.18	0.02*	0.25	0.02*
Euro	0.95	0.02*	0.41	0.02*	0.87	0.17	0.31	0.01**	0.38	0.02*
U.K.	0.32	0.02*	0.28	0.02*	0.49	0.02*	0.14	0.02*	0.22	0.02*
Canada	0.92	0.03*	0.13	0.02*	0.63	0.01**	0.30	0.02*	0.29	0.02*
Norway	0.49	0.02*	0.05*	0.02*	0.92	0.41	0.25	0.02*	0.33	0.02*
Sweden	0.85	0.01**	0.24	0.02*	0.45	0.01**	0.20	0.02*	0.29	0.02*
Switzerland	0.47	0.03*	0.36	0.02*	0.80	0.01**	0.01**	0.01**	0.07	0.01**
U.S.	0.09	0.01**	0.60	0.02*	0.29	0.02*	0.28	0.02*	0.04*	0.01**
Panel CADF-GLS unit root tests (Test statistics)										
CIPS	-2.1	-4.5**	-1.9	-4.8**	-1.8	-4.4**	-2.4*	-4.9**	-2.0	-4.5**
Fisher	58.4	478.3**	34.7	520.0**	50.1	547.8**	128.6**	418.4**	102.7**	469.4**
Inverse normal	-0.2	-17.1**	3.2	-18.3**	2.4	-18.4**	-5.0*	-16.1**	-2.5	-16.6**

See Note in Table B2 above.

Table B3: Unit root test statistics for the level and first differences of foreign variables in each country

Countries/ Regions	x_{it}^{*}	Δx_{it}^{*}	rp_{it}^{*}	Δrp_{it}^{*}	r_{it}^{*}	Δr_{it}^{*}	gov_{it}^{*}	Δgov_{it}^{*}
Individual ADF-GLS unit root tests (P-Values)								
China	0.87	0.02*	0.93	0.02*	0.06	0.02*	0.51	0.02*
Malaysia	0.10	0.02*	0.87	0.02*	0.05*	0.02*	0.59	0.02*
Indonesia	0.04*	0.02*	0.99	0.02*	0.05*	0.02*	0.56	0.02*
Philippines	0.11	0.02*	0.96	0.02*	0.08	0.02*	0.71	0.02*
Thailand	0.03*	0.02*	0.97	0.02*	0.05*	0.02*	0.75	0.02*
Vietnam	0.10	0.02*	0.98	0.02*	0.05*	0.02*	0.70	0.02*
Brunei	0.21	0.03*	0.86	0.02*	0.04*	0.02*	0.58	0.02*
India	0.13	0.02*	0.89	0.02*	0.05*	0.02*	0.83	0.02*
Pakistan	0.06	0.02*	0.73	0.02*	0.07	0.01**	0.85	0.02*
Sri Lanka	0.22	0.02*	0.66	0.02*	0.02*	0.02*	0.73	0.02*
Brazil	0.10	0.02*	0.99	0.01**	0.01**	0.03*	0.38	0.02*
Mexico	0.14	0.02*	0.33	0.01**	0.04*	0.02*	0.40	0.01**
Chile	0.24	0.02*	0.91	0.02*	0.01**	0.02*	0.40	0.02*
Turkey	0.46	0.01**	0.91	0.02*	0.12	0.36	0.92	0.02*
Argentina	0.22	0.02*	0.73	0.03*	0.02*	0.03*	0.44	0.03*
South Africa	0.34	0.02*	0.92	0.02*	0.03*	0.02*	0.93	0.02*
Japan	0.25	0.02*	0.93	0.01**	0.07	0.01**	0.84	0.02*
Korea	0.14	0.02*	0.95	0.02*	0.05*	0.02*	0.89	0.02*
Singapore	0.25	0.02*	0.91	0.02*	0.06	0.02*	0.72	0.02*
Australia	0.03*	0.02*	0.96	0.02*	0.12	0.01**	0.76	0.02*
New Zealand	0.03*	0.02*	0.25	0.01**	0.17	0.01**	0.85	0.40
Euro	0.19	0.02*	0.92	0.01**	0.03*	0.02*	0.66	0.02*
U.K.	0.32	0.02*	0.97	0.01**	0.13	0.25	0.91	0.02*
Canada	0.16	0.02*	0.61	0.01**	0.05*	0.01**	0.38	0.01**
Norway	0.98	0.01**	0.84	0.02*	0.17	0.46	0.83	0.01**
Sweden	0.34	0.01**	0.29	0.01**	0.13	0.41	0.88	0.02*
Switzerland	0.38	0.01**	0.84	0.02*	0.13	0.43	0.91	0.02*
U.S.	0.09	0.02*	0.57	0.02*	0.01**	0.02*	0.86	0.02*
Panel CADF-GLS unit root tests (Test statistics)								
CIPS	-3.1*	-5.3**	-2.2	-5.3**	-3.4**	-5.3**	-2.4	-5.0**
Fisher	112.1*	566.0**	37.1	532.3**	288.2**	694.4**	65.9	625.1**
Inverse normal	-4.3*	-19.8**	2.1	-19.3**	-12.7**	-22.7**	-1.0	-21.0**

See Note in Table B2.

Table B3 (continued)

Countries/ Regions	y^*_{it}	Δy^*_{it}	π^*_{it}	$\Delta\pi^*_{it}$	si^*_{it}	Δsi^*_{it}	e^*_{it}	Δe^*_{it}	oil_t	Δoil_t
Individual ADF-GLS unit root tests (P-Values)										
China	0.98	0.01**	0.16	0.02*	0.19	0.02*	0.61	0.01**		
Malaysia	0.93	0.05*	0.22	0.02*	0.58	0.02*	0.30	0.01**		
Indonesia	0.95	0.01**	0.25	0.02*	0.57	0.02*	0.23	0.01**		
Philippines	0.96	0.02*	0.26	0.02*	0.54	0.02*	0.59	0.01**		
Thailand	0.93	0.06	0.23	0.02*	0.54	0.02*	0.34	0.01**		
Vietnam	0.96	0.02*	0.26	0.02*	0.60	0.02*	0.25	0.01**		
Brunei	0.93	0.04*	0.06	0.03*	0.25	0.02*	0.40	0.01**		
India	0.90	0.02*	0.33	0.02*	0.16	0.02*	0.42	0.01**		
Pakistan	0.92	0.03*	0.32	0.02*	0.26	0.02*	0.34	0.02*		
Sri Lanka	0.92	0.03*	0.04*	0.02*	0.20	0.02*	0.31	0.02*		
Brazil	0.96	0.01**	0.28	0.02*	0.01**	0.02*	0.90	0.07		
Mexico	0.44	0.01**	0.34	0.02*	0.15	0.02*	0.29	0.02*		
Chile	0.98	0.01**	0.19	0.02*	0.02*	0.02*	0.62	0.02*		
Turkey	0.95	0.01**	0.44	0.01**	0.30	0.02*	0.14	0.01**		
Argentina	0.94	0.01**	0.16	0.02*	0.13	0.02*	0.06	0.02*		
South Africa	0.97	0.01**	0.42	0.02*	0.12	0.02*	0.29	0.01**		
Japan	0.77	0.12	0.24	0.02*	0.33	0.02*	0.70	0.01**		
Korea	0.91	0.10	0.25	0.02*	0.61	0.02*	0.82	0.01**		
Singapore	0.90	0.02*	0.03*	0.02*	0.28	0.02*	0.50	0.01**		
Australia	0.94	0.04*	0.31	0.02*	0.69	0.01**	0.37	0.01**		
New Zealand	0.94	0.04*	0.34	0.02*	0.39	0.02*	0.31	0.01**		
Euro	0.96	0.01**	0.40	0.02*	0.18	0.02*	0.70	0.01**		
U.K.	0.96	0.01**	0.48	0.02*	0.40	0.02*	0.19	0.01**		
Canada	0.26	0.01**	0.41	0.02*	0.56	0.01**	0.92	0.01**		
Norway	0.97	0.01**	0.19	0.01**	0.48	0.01**	0.08	0.02*		
Sweden	0.98	0.01**	0.19	0.01**	0.44	0.01**	0.14	0.02*		
Switzerland	0.97	0.01**	0.22	0.01**	0.42	0.02*	0.15	0.01**		
U.S.	0.98	0.01**	0.44	0.02*	0.45	0.02*	0.82	0.01**	0.90	0.02*
Panel CADF-GLS unit root tests (Test statistics)										
CIPS	0.1	-2.8**	-1.8	-5.4**	-1.8	-4.9**	-2.5	-4.7**		
Fisher	15.9	235.1**	60.9	660.5**	94.7*	537.0**	42.7	493.8**		
Inverse normal	11.0	-9.1**	-1.6	-22.5**	-2.4	-19.7**	0.4	-18.7**		

See Note in Table B2 above.

Table B4: Unit root test statistics for the level and first differences of differential variables in each country

Countries/ Regions	$x^*_{it} - x_{it}$	$\Delta(x^*_{it} - x_{it})$	$rp^*_{it} - rp_{it}$	$\Delta(rp^*_{it} - rp_{it})$	$r^*_{it} - r_{it}$	$\Delta(r^*_{it} - r_{it})$
Individual ADF-GLS unit root tests (P-Values)						
China	0.89	0.01**	0.84	0.02*	0.03*	0.01**
Malaysia	0.48	0.02*	0.88	0.02*	0.01**	0.02*
Indonesia	0.01**	0.02*	0.36	0.02*	0.72	0.75
Philippines	0.01**	0.01**	0.28	0.02*	0.01**	0.02*
Thailand	0.01**	0.77	0.89	0.02*	0.01**	0.02*
Vietnam	0.52	0.01**	0.66	0.03*	0.79	0.06
Brunei	0.10	0.67	0.59	0.02*	0.11	0.99
India	0.03*	0.02*	0.72	0.02*	0.01**	0.02*
Pakistan	0.29	0.03*	0.32	0.02*	0.01**	0.02*
Sri Lanka	0.41	0.01**	0.05*	0.02*	0.01**	0.03*
Brazil	0.51	0.02*	0.71	0.04*	0.05*	0.02*
Mexico	0.70	0.01**	0.27	0.02*	0.01**	0.02*
Chile	0.13	0.01**	0.26	0.22	0.01**	0.02*
Turkey	0.90	0.01**	0.34	0.02*	0.01**	0.02*
Argentina	0.02*	0.02*	0.40	0.02*	0.01**	0.02*
South Africa	0.44	0.02*	0.48	0.02*	0.02*	0.02*
Japan	0.83	0.01**	0.56	0.02*	0.01**	0.03*
Korea	0.97	0.01**	0.35	0.02*	0.07	0.01**
Singapore	0.14	0.01**	0.58	0.01**	0.01**	0.23
Australia	0.62	0.02*	0.40	0.02*	0.02*	0.03*
New Zealand	0.75	0.02*	0.54	0.02*	0.02*	0.83
Euro	0.65	0.02*	0.35	0.02*	0.02*	0.02*
U.K.	0.08	0.02*	0.10	0.01**	0.01**	0.03*
Canada	0.51	0.01**	0.67	0.01**	0.01**	0.02*
Norway	0.32	0.02*	0.90	0.01**	0.01**	0.03*
Sweden	0.91	0.01**	0.87	0.01**	0.02*	0.03*
Switzerland	0.33	0.02*	0.30	0.02*	0.01**	0.03*
U.S.	0.23	0.01**	0.77	0.02*	0.01**	0.02*
Panel CADF-GLS unit root tests (Test statistics)						
CIPS	-2.6	-5.5**	-2.1	-4.8**	-3.0**	-4.9**
Fisher	46.1	397.3**	40.2	521.1**	181.5**	567.1**
Inverse normal	-1.1	-17.6**	2.1	-19.3**	-7.4**	-18.0**

See Note in Table B2.

Table B4 (continued)

Countries/ Regions	$gov^*_{it} - gov_{it}$	$\Delta(gov^*_{it} - gov_{it})$	$y^*_{it} - y_{it}$	$\Delta(y^*_{it} - y_{it})$
Individual ADF-GLS unit root tests (P-Values)				
China	0.41	0.02*	0.11	0.04*
Malaysia	0.45	0.02*	0.92	0.74
Indonesia	0.21	0.03*	0.99	0.01**
Philippines	0.15	0.01**	0.99	0.01**
Thailand	0.11	0.01**	0.92	0.02*
Vietnam	0.01**	0.01**	0.11	0.69
Brunei	0.93	0.01**	0.93	0.02*
India	0.02*	0.01**	0.94	0.01**
Pakistan	0.16	0.02*	0.87	0.01**
Sri Lanka	0.43	0.03*	0.88	0.15
Brazil	0.20	0.02*	0.06	0.02*
Mexico	0.39	0.02*	0.98	0.03*
Chile	0.48	0.27	0.95	0.03*
Turkey	0.32	0.02*	0.61	0.02*
Argentina	0.38	0.01**	0.92	0.04*
South Africa	0.34	0.01**	0.09	0.59
Japan	0.63	0.01**	0.84	0.82
Korea	0.14	0.02*	0.76	0.01**
Singapore	0.20	0.02*	0.70	0.02*
Australia	0.22	0.02*	0.77	0.87
New Zealand	0.08	0.01**	0.74	0.01**
Euro	0.35	0.02*	0.21	0.10
U.K.	0.18	0.01**	0.40	0.01**
Canada	0.28	0.01**	0.88	0.01**
Norway	0.50	0.02*	0.94	0.22
Sweden	0.08	0.02*	0.96	0.01**
Switzerland	0.29	0.01**	0.94	0.02*
U.S.	0.63	0.01**	0.97	0.68
Panel CADF-GLS unit root tests (Test statistics)				
CIPS	-1.5	-4.5**	-1.6	-3.8**
Fisher	47.3	434.5**	34.1	189.0**
Inverse normal	-0.2	-16.5**	0.5	-10.5**

See Note in Table B2 above.

Table B5: Panel cointegration tests and DOLS cointegrating vectors

These tables show the results of Pedroni cointegration tests for the long-run relationships (2)-(4) in Section 4.3. The numbers in the columns named Panel PP, Group PP, Panel ADF, and Group ADF indicate probability values of the tests under the null hypothesis of no cointegration. If there is enough evidence in favor of cointegrating relationships between variables at the 5% significance level, we compute cointegrating vectors using DOLS in the case of developed country groups and we use cointegrating vectors suggested by economic theory in the case of developing country groups (due to the short spans of the data set in these countries). This is shown in the last column of the tables.

Table B5.1: Probability values associated with cointegration tests for the Fisher Equation

si_{it}, π_{it}	Panel PP	Group PP	Panel ADF	Group ADF	Cointegrating vector
European	0.09	0.21	0.17	0.34	–
Other-developed	0.27	0.63	0.28	0.65	–
Developing-Asia	–	–	–	–	–
Other-developing	0.00	0.00	0.00	0.00	$(1, -1)$

Note: 1) The figures in parentheses are the Panel DOLS estimates of cointegrating vectors.

2) The number in the table stands for the P-value for the tests under the null hypothesis of no cointegration.

3) The lag length in ADF-type regression of Pedroni's tests is selected by BIC with a maximum of 4 lags.

4) The critical value is determined by the standard normal distribution.

Table B5.2: Probability values associated with cointegration tests for the Output Convergence

y_{it}, y^*_{it}	Panel PP	Group PP	Panel ADF	Group ADF	Cointegrating vector
European	0.92	0.98	0.80	0.48	–
Other-developed	0.55	0.60	0.54	0.14	–
Developing-Asia	0.78	0.70	0.83	0.22	–
Other-developing	0.18	0.26	0.02	0.12	–

Note: See Table B5.1

Table B5.3: Probability values associated with cointegration tests for the UIP Condition

si_{it}, si^*_{it}	Panel PP	Group PP	Panel ADF	Group ADF	Cointegrating vector
European	0.00	0.00	0.00	0.00	$(1, -0.641)$
Other-developed	0.01	0.02	0.00	0.00	$(1, -0.940)$
Developing-Asia	0.00	0.00	0.00	0.00	$(1, -1)$
Other-developing	0.01	0.11	0.00	0.00	–

Note: See Table B5.1.

Table B6: The "general-to-particular" mode of search for the irreducible cointegrating relation

These tables show the test results of the general-to-particular mode of search. First, we start with the collection of variables which includes the real exchange rate and the 6 fundamentals. Then we drop each variable one at a time and conduct Pedroni cointegration tests to see whether the remaining variables are cointegrated. If the remaining variables are still cointegrated after one variable is dropped, this indicates that the omitted variable is redundant. However, if the remaining variables are not cointegrated, the omitted variable is important for the cointegrating relation and cannot be dropped.

Table B6.1: Cointegration tests for a relationship starting from the real exchange rate and 6 fundamentals

The collection of variables	Panel PP	Group PP	Panel ADF	Group ADF
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}),$ $(y^*_{it} - y_{it}), tt_{it}, open_{it}$	0.00	0.00	0.00	0.03
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}),$ $(y^*_{it} - y_{it}), tt_{it}$	0.99	0.02	0.99	0.00
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}),$ $(y^*_{it} - y_{it}), open_{it}$	0.00	0.19	0.00	0.58
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}),$ $tt_{it}, open_{it}$	0.00	0.09	0.00	0.07
$q_{it}, (x^*_{it} - x_{it}), (rp^*_{it} - rp_{it}), (y^*_{it} - y_{it}),$ $tt_{it}, open_{it}$	0.00	0.35	0.00	0.11
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}),$ $tt_{it}, open_{it}$	**0.00**	**0.01**	**0.00**	**0.00**
$q_{it}, (rp^*_{it} - rp_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}),$ $tt_{it}, open_{it}$	0.00	0.06	0.00	0.04

Note: 1) The number in the table stands for the P-value for the tests under the null hypothesis of no cointegration.
2) The lag length in ADF-type regression of Pedroni's tests is selected by BIC with a maximum of 4 lags.
3) The critical value is determined by the standard normal distribution.

The first collection (the real exchange rate and its 6 fundamentals) is the initial cointegrating group suggested by Pedroni's test. The criterion decision that we use is that a variable will be viewed as an irrelevant variable if after dropping it, all four tests are still statistically significant at 5% level. Table B6.1 shows that only the relative price of nontraded goods $(rp^*_{it} - rp_{it})$ is irrelevant to the cointegrating relationship and can be omitted. Then we check whether the set of variables obtained after omitting the relative price contains an irrelevant variable. The test results are as follows.

Table B6.2: Cointegration tests for a relationship starting from the real exchange rate and 5 fundamentals

The collection of variables	Panel PP	Group PP	Panel ADF	Group ADF
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), tt_{it}, open_{it}$	0.00	0.01	0.00	0.00
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), tt_{it}$	0.99	0.17	0.99	0.00
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), open_{it}$	**0.00**	**0.03**	**0.00**	**0.02**
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), tt_{it}, open_{it}$	0.00	0.23	0.00	0.08
$q_{it}, (x^*_{it} - x_{it}), (y^*_{it} - y_{it}), tt_{it}, open_{it}$	0.00	0.37	0.00	0.47
$q_{it}, (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), tt_{it}, open_{it}$	0.00	0.10	0.00	0.06

Note: See Table B6.1.

In Table B6.2, it is clear that the irrelevant variable is the terms of trade, tt_{it} for this set of variables. We then continue the process on the new collection that excludes both the relative price and the terms of trade. We obtain the following results.

Table B6.3: Cointegration tests for a relationship starting from the real exchange rate and 4 fundamentals

The collection of variables	Panel PP	Group PP	Panel ADF	Group ADF
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), open_{it}$	0.00	0.03	0.00	0.02
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it})$	0.99	0.14	0.99	0.19
$q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), open_{it}$	0.00	0.06	0.00	0.05
$q_{it}, (x^*_{it} - x_{it}), (y^*_{it} - y_{it}), open_{it}$	0.00	0.26	0.00	0.40
$q_{it}, (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it}), open_{it}$	0.00	0.07	0.00	0.08

Note: See Table B6.1.

As before, the first row is the intial group of variables obtained from the previous test. Now, it is easy to see that no variable can be omitted without losing the cointegration property at the 5% significance level. Therefore, the group of variables comprising $q_{it}, (x^*_{it} - x_{it}), (gov^*_{it} - gov_{it}), (y^*_{it} - y_{it})$, and $open_{it}$ is " the irreducible cointegrating relation".

Table B7: F statistics for testing the weak exogeneity of the country/region-specific foreign variables

	Foreign Variables					
Country Groups	x^*_{it}	y^*_{it}	gov^*_{it}	si^*_{it}	inf^*_{it}	oil_t
European	1.323	0.025	1.907	2.131	0.243	1.735
	[0.269]	[0.975]	[0.152]	[0.122]	[0.785]	[0.181]
Other-developed	0.820	0.526	0.732	0.842	2.835	1.860
(excluding US)	[0.442]	[0.592]	[0.482]	[0.062]	[0.062]	[0.159]
Developing-Asia	1.254	2.570	0.658	1.773	2.303	1.315
	[0.292]	[0.056]	[0.579]	[0.154]	[0.078]	[0.270]
Other-developing	0.291	1.669	0.411	2.457	3.623*	2.642
	[0.748]	[0.192]	[0.664]	[0.090]	[0.030]	[0.075]

Note: The figures in square brackets are estimated probability values.

* denotes statistical significance at the 5% level or less.

Table B8: Impulse responses of the shocks to exchange rate determinants on real exchange rates

Country Groups	**Shocks** Productivity Improvement	Trade Liberalization	Contractionary Monetary Policy	Government Spending
Expected response	Appreciate	Depreciate	Appreciate	Appreciate
European				
Relationship	Depreciate	Depreciate	No response	Appreciate
Time period	from beginning	until 3rd yr	-	from beginning
Persistent	Yes	No	-	Yes
Other-developed				
Relationship	Depreciate	Depreciate	Appreciate	Depreciate
Time period	until 1st yr	from beginning	until 5th yr	from 1st yr
Persistent	No	Yes	No	Yes
Developing-Asia				
Relationship	No response	Depreciate	No response	Appreciate
Time period	-	from beginning	-	from 6th yr
Persistent	-	Yes	-	Yes
Other-developing				
Relationship	Appreciate	Depreciate	Appreciate	Appreciate
Time period	from 3rd yr	from beginning	until 2nd yr	from 1st yr
Persistent	Yes	Yes	No	Yes

Table B9: Forecast Error Variance Decomposition for real exchange rates, reported at peak during 9 years following the shocks (%)

	Shocks				
Country Groups	Productivity Improvement	Trade Liberalization	Contractionary Monetary Policy	Government Spending	Total
European	7.89	15.08	4.85	12.34	40.16
	[19.65]	[37.55]	[12.06]	[30.74]	[100.00]
Other-developed	3.95	55.64	4.66	6.62	70.88
	[5.57]	[78.50]	[6.58]	[9.35]	[100.00]
Developing-Asia	1.55	26.95	4.86	1.39	34.75
	[4.45]	[77.56]	[14.00]	[4.00]	[100.00]
Other-developing	6.73	6.95	6.76	15.47	35.91
	[18.75]	[19.34]	[18.83]	[43.07]	[100.00]

Note: The numbers in square brackets represent the percentage of variance explained by each shock to the total variance explained by the four shocks in each country group.

C Figure Appendix

Figure C1: Time plots of some series during Latin American Debt Crisis

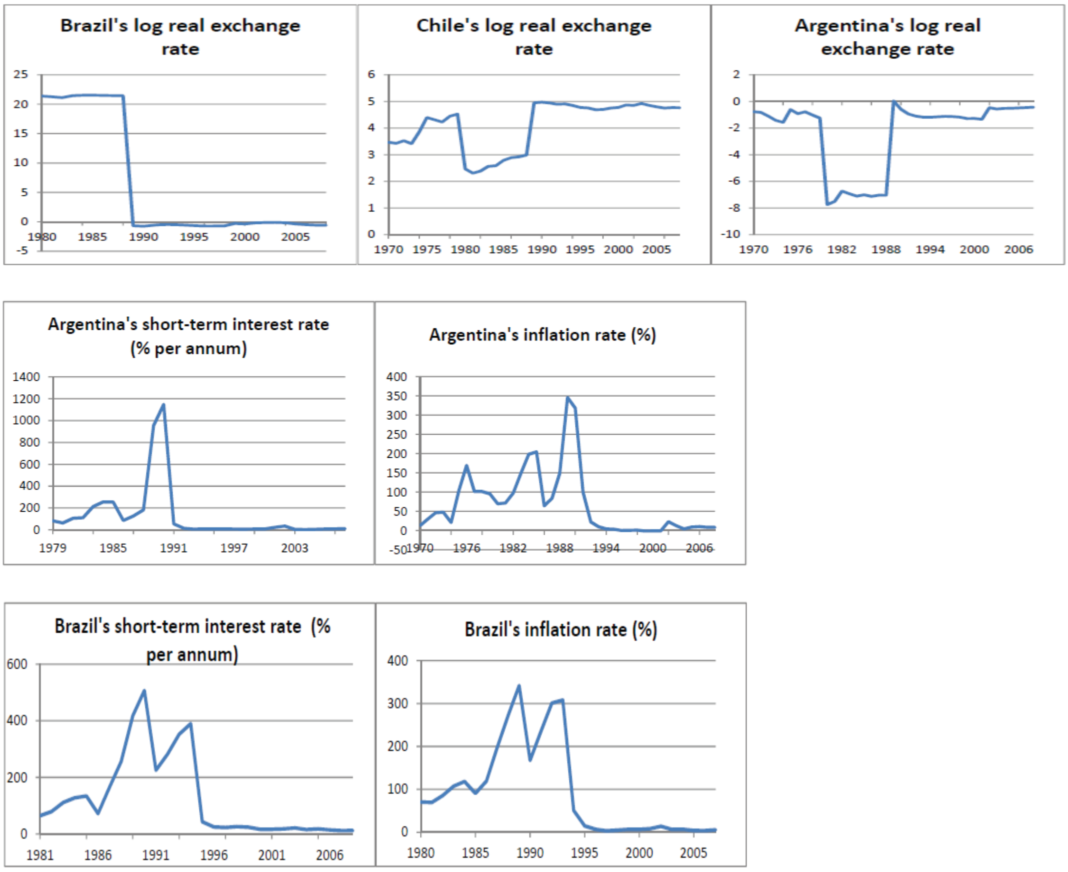

Figure C2: Impulse responses of the European group to a productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

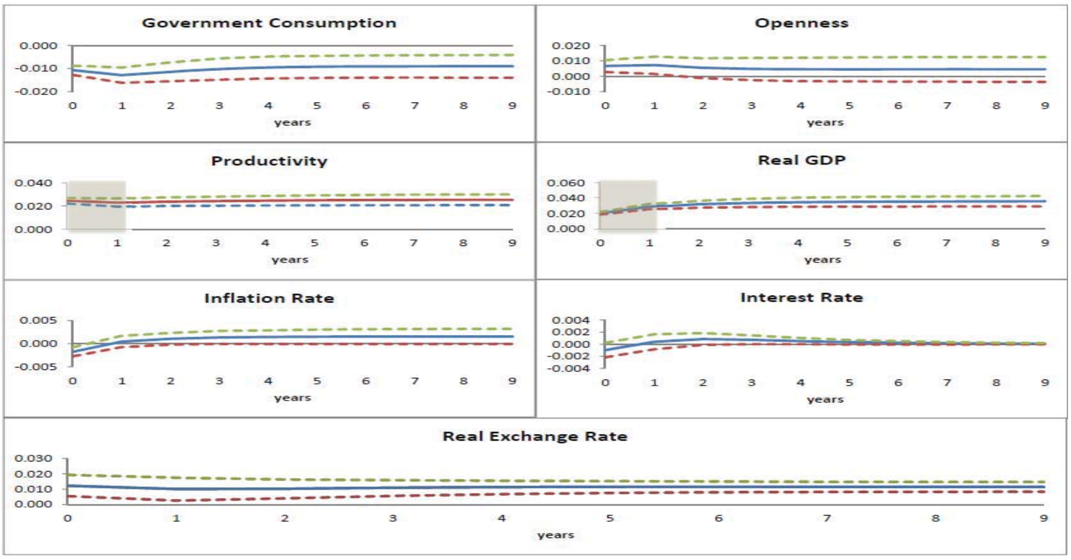

Figure C3: Impulse responses of the other-developed country group to a productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

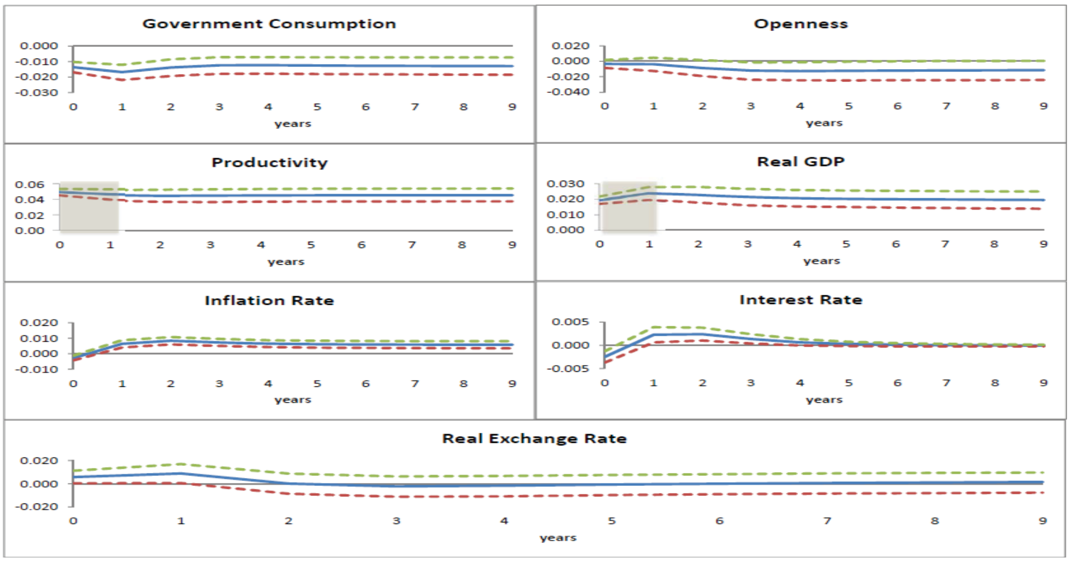

Figure C4: Impulse responses of the developing-Asia group to a productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

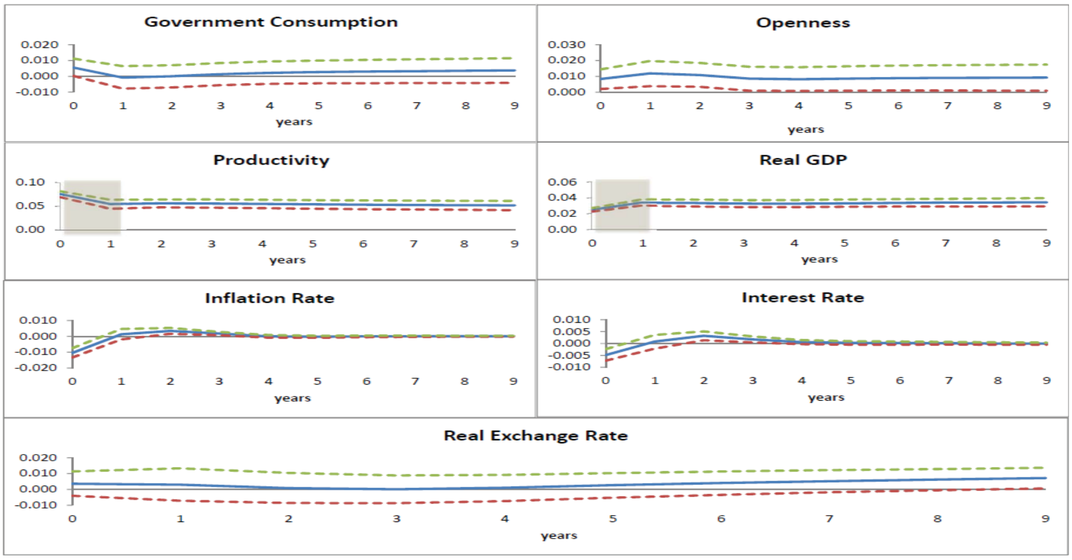

Figure C5: Impulse responses of the non Asian developing-country group to a productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

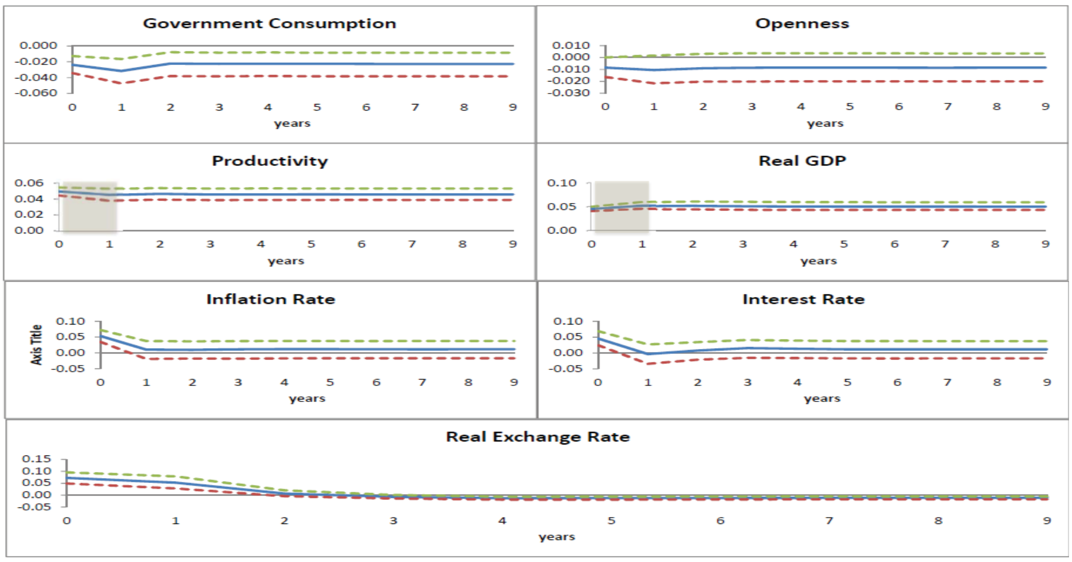

Figure C6: Impulse responses of the European group to a trade-liberalization shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

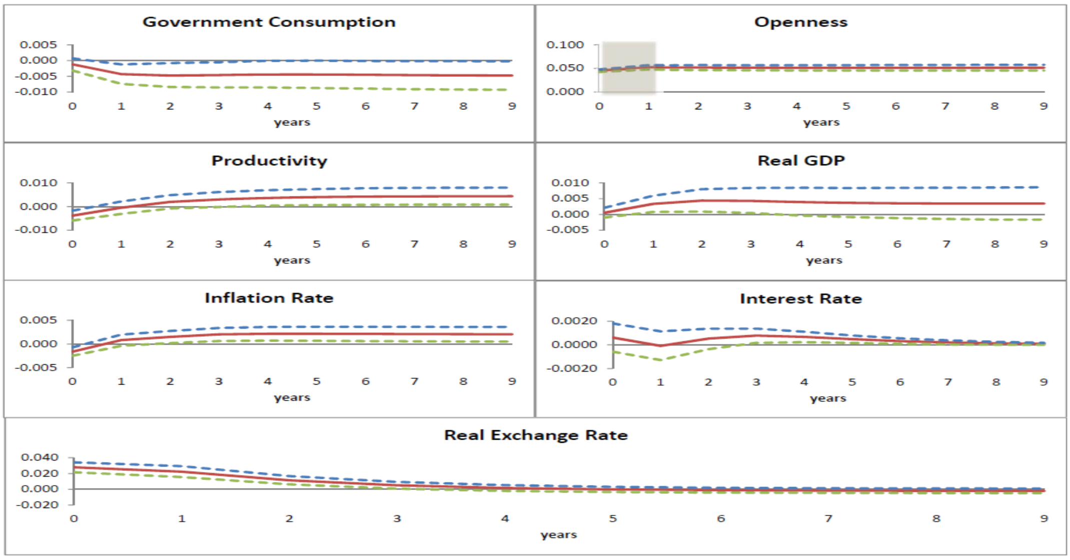

Figure C7: Impulse responses of the other-developed country group to a trade-liberalization shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

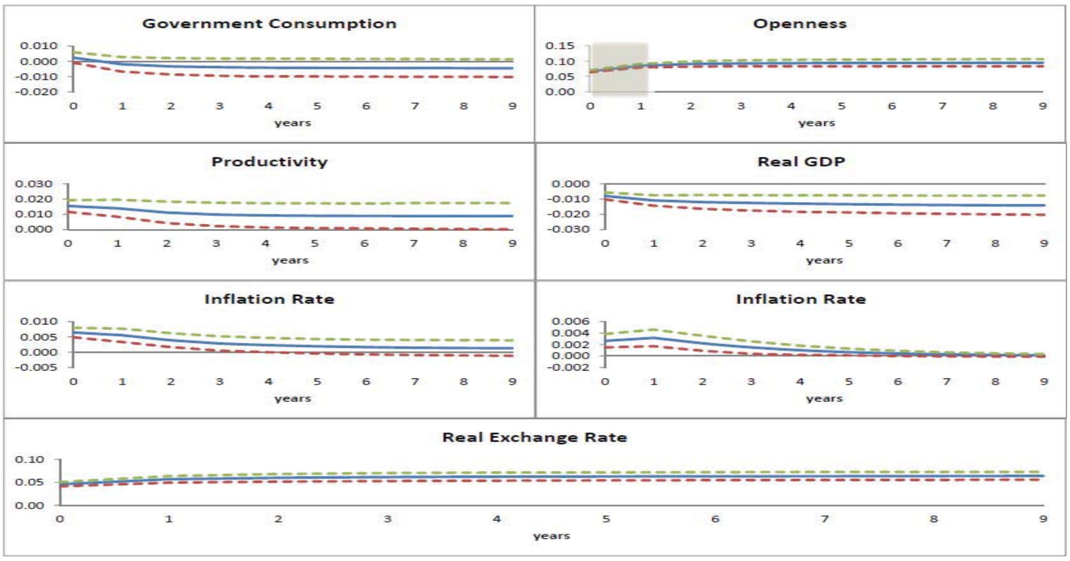

Figure C8: Impulse responses of the developing-Asia group to a trade-liberalization shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

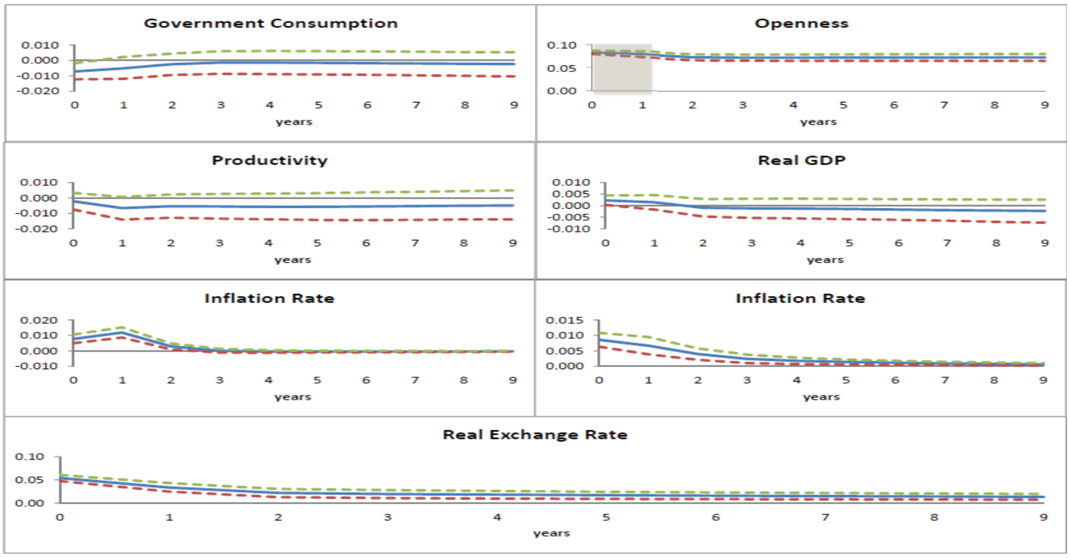

Figure C9: Impulse responses of the non Asian developing-country group to a trade-liberalization shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

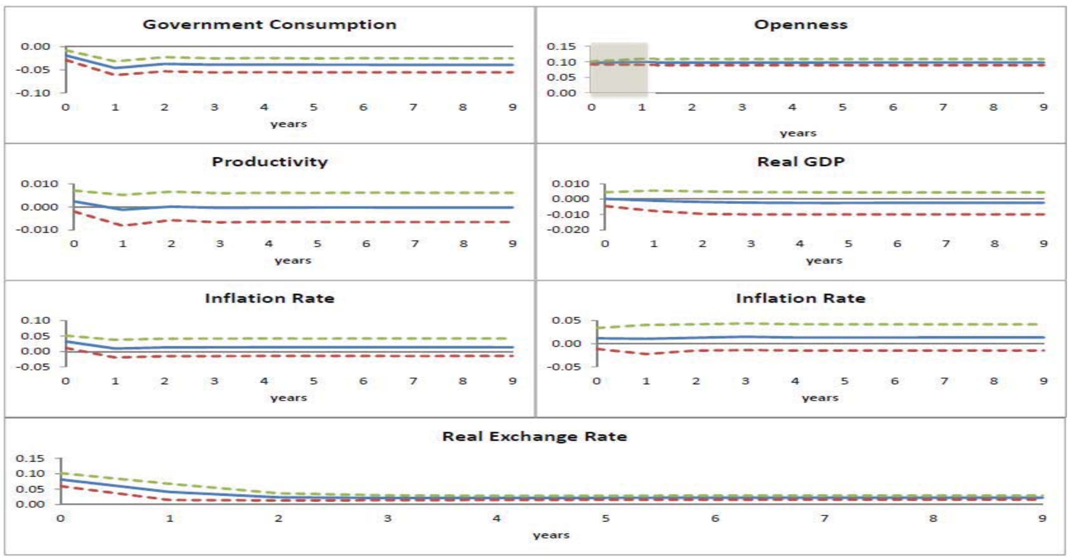

Figure C10: Impulse responses of the European group to a contractionary monetary policy shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

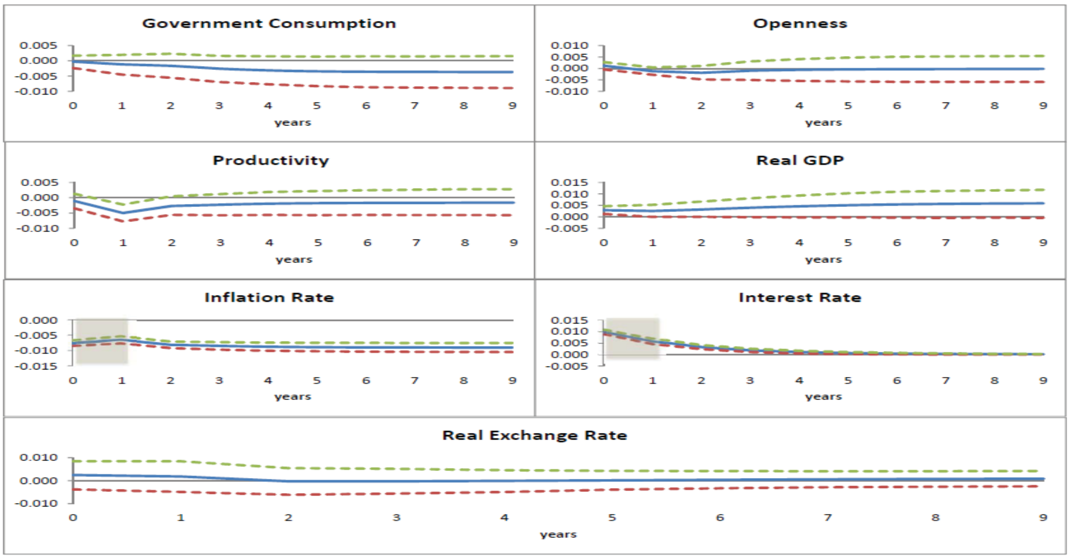

Figure C11: Impulse responses of the other-developed country group to a contractionary monetary policy shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

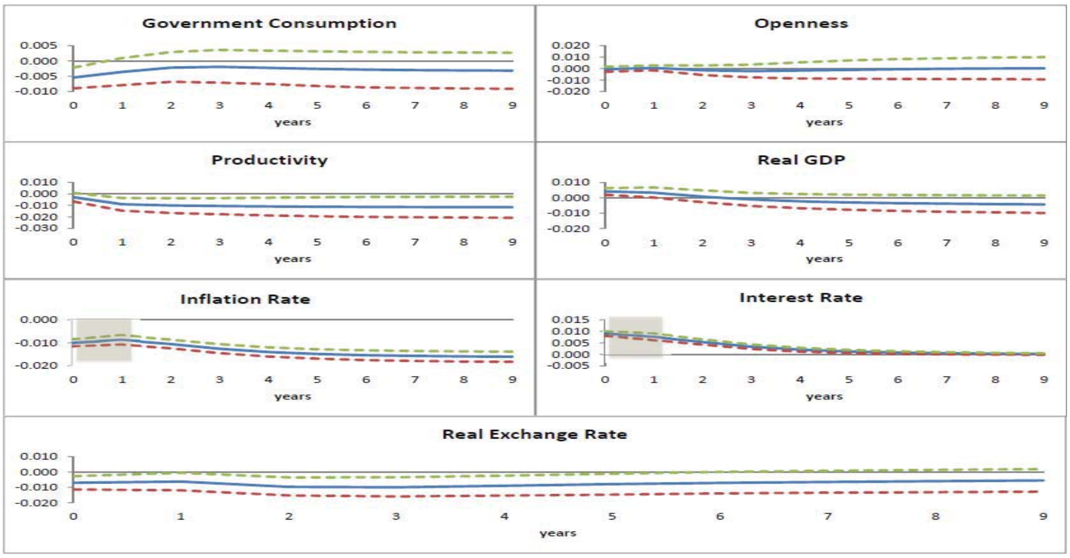

Figure C12: Impulse responses of the developing-Asia group to a contractionary monetary policy shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

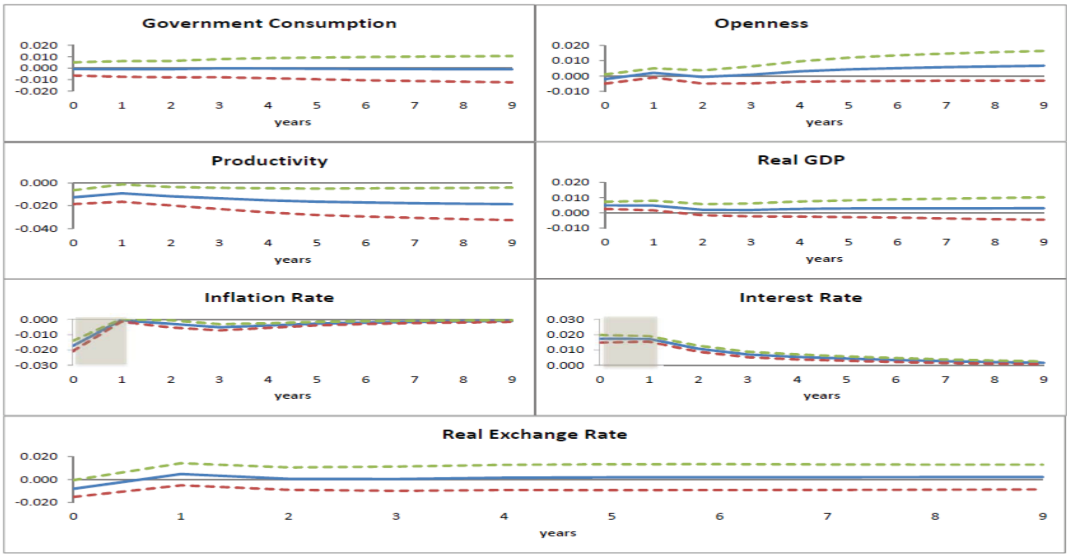

Figure C13: Impulse responses of the non Asian developing-country group to a contractionary monetary policy shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

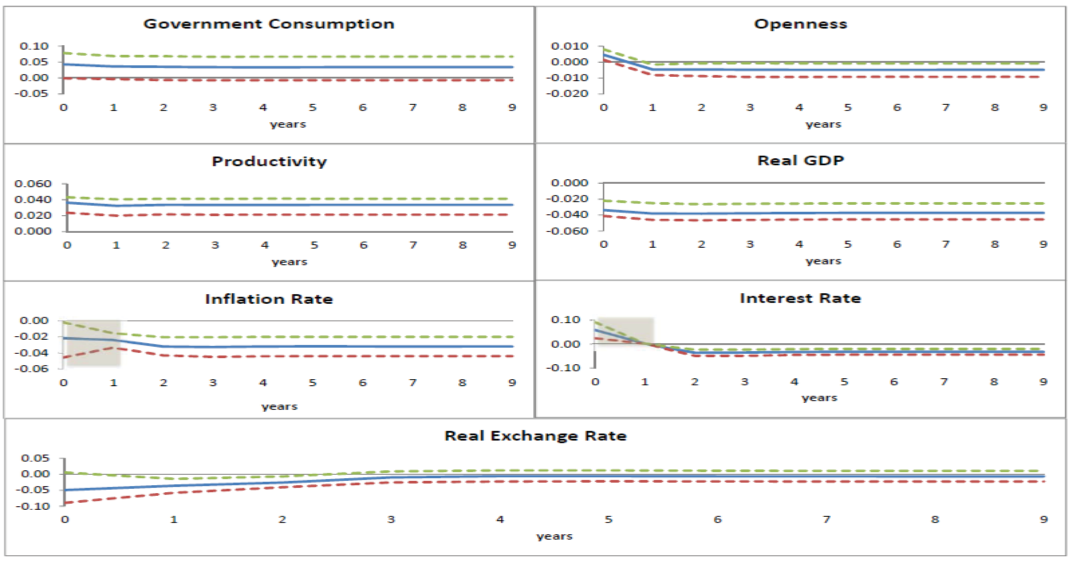

Figure C14: Impulse responses of the European group to a government spending shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

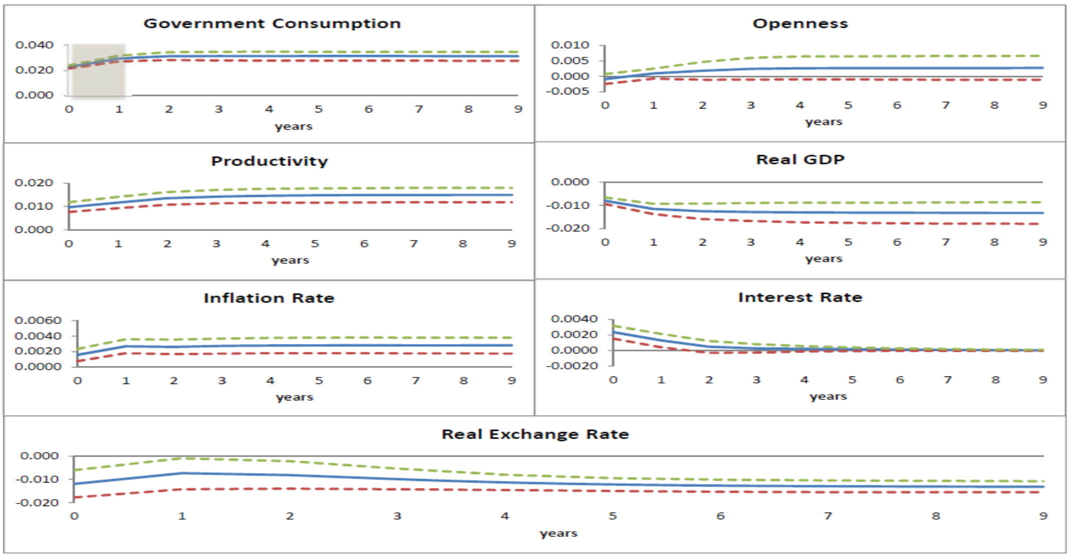

Figure C15: Impulse responses of the other-developed country group to a government spending shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

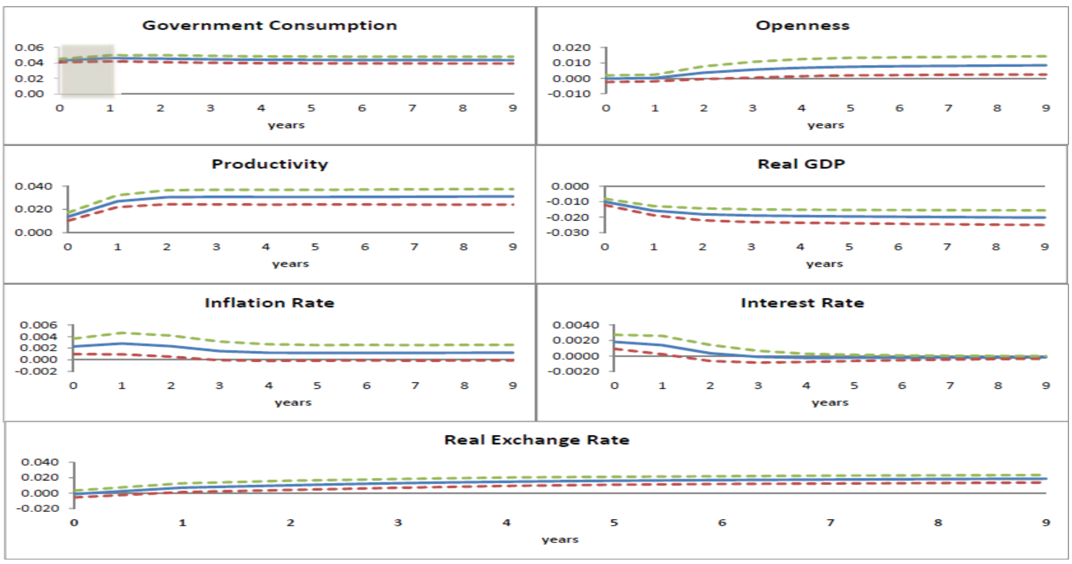

Figure C16: Impulse responses of the developing-Asia country group to a government spending shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.

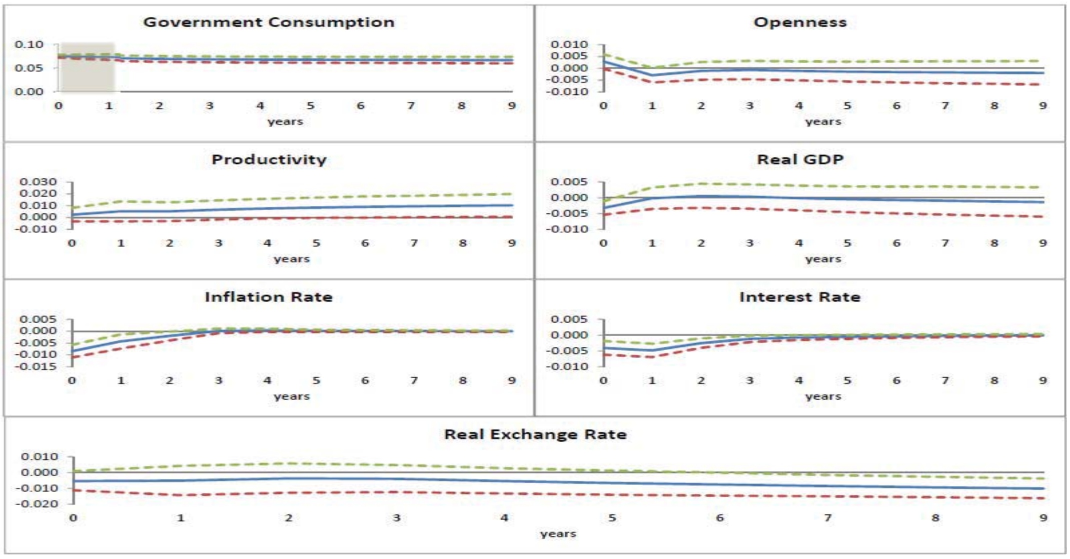

Figure C17: Impulse responses of the non Asian developing-country group to a government spending shock of one standard deviation in size, using sign restrictions with the penalty function. The shaded areas indicate the responses restricted by sign restrictions.